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Aspects of Business Cycle Transmission in Europe

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Contents

Introduction

The reader of the business cycles literature might be puzzled by the contrast between intuitive definitions that seem plain at first sight – e.g. "short-term fluctuations in the level of economic activity, relative to the long-term trend in output"[1] or "periodic swings in the pace of national economic activity, characterized by alternating expansion and contraction phases"[2] – and the technical difficulties of empirical identifications of business cycles. There is indeed a gap between what we could define as a *broad* definition of business cycles and a more rigorous and precise one. Even though the business cycles literature has been developed for decades[3] there are still wide-ranging debates concerning the exact definition of the business cycles and the way to identify them (see e.g. Eurostat, 2004). So far, the literature has not reached any consensus on the best way to identify a business cycle. Indeed, it does not pre-exist the analysis. There is a multiplicity of business cycles and they largely depend upon what the researcher is looking for. One of the most consensual definition is the one of Burns and Mitchell (1946) who define a business cycle as elements of macroeconomic series having a periodicity contained between one and ten or twelve years[4]. Another approach is to consider the cycle as the remaining, stationary part, once the trend has been extracted from a series. This decomposition into a trend and a cycle reflects the idea that fluctuations are independent from the long-run trend of the economy. In this view, cycles might be seen as accidents, in that they are simply deviations from the equilibrium. Short term movements are completely separated from long term ones. As we will see in the second part of the present work, these two definitions may be compatible but are not strictly equivalent.

The first type of cycles we use is based on the definition of Burns & Mitchell. We will not talk about classifications such as the Kondratieff cycle (with a periodicity of 50-60 years), the Juglar cycle (8 years on average) or the Kitchin cycle (3.5). The approach based on frequencies generalizes these classifications since time series are seen as decompositions of an infinity of elements, each having a particular periodicity. This is the type of cycles that will be used in the second part. In the two last parts a trend/cycle decomposition

[1] among other definitions found on the Internet: lms.thomsonelearning.com/hbcp/glossary/glossary.taf

[2] http://www.garlic.com/~lynn/paygloss.htm

[3] e.g. Burns & Mitchell (1946) or Hicks (1950)

[4] "...a cycle consists of expansions occurring at about the same time in many economic activities, followed by similarly general recessions, contractions, and revivals which merge into the expansion phase of the next cycle (...) In duration business cycles vary from more than one year to ten or twelve years..." (Burns and Mitchell, 1946, p 3).

has been chosen.

Many theories have emerged to understand how and why business cycles appear and endure. There has been an evolution in theory. The dominant thought after World War II was that business cycles were self-sustaining. For instance, this was the idea of the Keynesian 'Samuelson oscillator' model. Hicks (1950) used a model of the same type and introduced nonlinearity in order to improve the ability of the system to fluctuate endogenously. One could also quote the model of Goodwin (1967), who showed that for some values of the model parameters, cycles in production and employment were appearing endogenously. This model was developed later by Jarsulic (1986), who extended it in order to take into account the idea of 'limit cycles' and attractors. We could also mention developments in recent years linked to the chaos theory.

The problem of these models, in particular those of the accelerator-multiplier type, is that the cycles produced cannot replicate the erratic movements observed in reality. This may be one of the elements explaining why the dominant notion of business cycles has moved towards an explanation of the Wicksellian type. In this perspective, fluctuations are the effect of shocks hitting the economy. Wicksell used the following metaphor (quoted by Frisch, 1933): "If you hit a wooden rocking-horse with a club, the movement of the horse will be very different from that of the club.". Using this idea, Frisch was one of the first to model separately a propagation mechanism –his model was creating dampening fluctuations– and the impulse mechanism –composed of cumulative stochastic shocks. As asserted by Abraham-Frois (1995), Frisch has had a direct influence on rational expectations business cycles models. The debate at this point was centered on the nature of shocks that create cycles. For Lucas (1973), the impulse of cycles comes from monetary shocks. At the opposite, Real Business Cycles (RBC) emphasize on real shocks –technological or budgetary. RBC models have been central in the literature during the past years, and these models were able to produce many 'stylized facts' of business cycles.

The purpose here is not to take a stand on whether cycles result from exogenous shocks, or whether fluctuations are endogenous to the economy. We will rather observe existing cycles and compare them. In particular, we will try to examine the influence of cycles on other economic entities, for instance other countries' cycles or other macroeconomic variables within the same country.

Zarnowitz (1992) has described some stylized facts about business cycles. For example, sectors producing manufactured goods are more cyclical than services. Similarly, private investment is more cyclical than consumption. But the most noticeable facts are certainly that (1) business cycles have quite often an international dimension and (2) that business cycles affect the whole economy: production, employment, investment and all the main macroeconomic aggregates –see the definition of Burns & Mitchell above. The following work will examine these two facts, focusing essentially on the first point (parts one to three) and skimming over the second (part four).

We will examine different aspects of the transmission of business cycles. Transmission in the sense that the business cycle influences other macroeconomic variables. A particular attention will be dedicated to the influence of a cycle in one particular geographical area onto cycles of other areas. For example, we will look at the mutual influence of cycles in Europe in the first part. In the second, we will try to understand what is the channel through which business cycles are transmitted from one country to another. International trade is believed to be such a channel. The third part will emphasize on the relation between a Euro area cycle and the cycle of an 'external' country.

Another kind of transmission will be studied in the last part. We will have a brief view on the influence of business cycles on other macroeconomic variables. We will take one specific geographic area –the Euro area– in order to estimate the so-called natural rate of interest. It is assumed in this model that the business cycle has an influence on inflation and that there is a mutual influence between the cycle and the natural real interest rate. To the extent that the natural rate of interest is taken into account by the ECB, we could say that business cycles also have an influence on monetary policy.

The first three parts are centered on the influence of cycles between economies. In a sense, the first and the third paper try to answer the question 'have business cycles become more similar?', while it is asked in the second, '*why* have they become more similar?'. The last paper examine the influence of cycles *within* an economy.

We focus on countries of the European Union and of the Euro zone. When needed, we use other countries such as the US or Japan in order to make comparisons.

Even if it is not the main issue of this thesis, Optimal Currency Area theory will implicitly guide our reflection in the following parts. The main idea is that coordinated business cycles is a necessary condition for the area to be optimal for a single currency. We will come back on this later on. A problematic feature of this theory is that it is rather vague as to the degree and the type of business cycles coordination necessary to get optimality. This is why we will consider different measures of business cycles. We will try to keep in mind the arguments of the OCA theory while drawing our conclusions.

The first chapter is concerned with the similarity of business cycles in the European Union. It compares the cycles across countries before and after the launching of the European Monetary System. This comparison is based on two criteria: average shapes and timing of cycles. For this purpose, we use the 'classical cycles' approach. It consists in using a particular procedure to find the peaks and troughs in a series –here, industrial production. Business cycles are defined as the part contained between two peaks or two troughs. The main result is that there is a core group of European countries for which the links are quite strong and have increased in time.

Starting from the observation made in the previous chapter that business cycles tend to be coordinated across European countries –more strongly for some than for others–, the second chapter tries to identify a channel through which this coordination can occur. Trade is often cited as an important factor in business cycle transmission. Industrial countries tend to exchange similar goods ('Intra-Industry Trade' – *IIT*). Consequently, these economies should be affected by similar shocks and their business cycles should be synchronized. This part is composed of two sub-chapters. In the first one, the definition of business cycles is based on periodicity considerations. We introduce here the frequency domain and the tools used allow in particular to distinguish cycles that exhibit comovements from cycles that are synchronized[5]. The main result is that there is some evidence of a link between the trade structure and cycles similarity. It seems that IIT is correlated with cycles that comove and are synchronized at the same time, i.e. cycles that might be the fact of common supply shocks. Conversly, there seems to be a negative influence on cycles that comove with a time delay. In the second sub-chapter, we take a more conventional measure for business cycles, and we put emphasis on the trade structure. A distinction

[5]The former concept refers to series that exhibit movements of similar amplitude and frequency, but that are not necessarily affected by a shock at the same time. In contrast, the latter concept refers to series that react at the same time to a given shock, but for which the response in amplitude or frequency can be different.

is made between 'vertical' and 'horizontal' IIT. In other words, goods of the same type are differenciated between those that are substitutable and those that exhibit large differences in their unit prices. The estimations show that business cycles correlations are more strongly (positively) linked to vertical IIT than to horizontal IIT. An interpretation of this result could be that VIIT concerns industries that are subject to common supply shocks and that are hit by foreign demand shocks at the same time. In other words, an idiosyncratic demand shock will propagate abroad if the goods are not substitutable. If they are substitutable –as it is the case for HIIT– consumers might prefer to buy at home and demand shocks will not be propagated.

The third part looks more specifically at the UK. While studying business cycle relations, authors often notice that the UK cycle was quite independent from the other European cycles. It is to some extent quite similar to the US one, and it lies somehow between the two continents. This was one of the argument against the entrance of the UK in the EMU. At the same time, some authors have pointed out that the UK cycle is getting closer to the 'continental' one. In this paper, we take the question from a slightly different perspective. Basically, we will try to see if the coherence of a core group of EMU countries is affected when we include the UK. To this end, we address the issue of the desirability of the entrance of the UK from the insiders point of view. The technique used in this part is based upon the Kalman filter. The definition of business cycles is different than in the two previous parts. It corresponds to the second definition, i.e. the trend/cycles decomposition. Cycles are not defined in terms of frequencies but are considered as unobserved components. The researcher does not observe directly the cycles but makes some assumptions about the way they behave in time. The Kalman filter will provide a way to extract the cycles and eventually to distinguish between idiosyncratic and common cycles. Indeed, we put forward the hypothesis that members of a particular group of countries share a common cycle. The purpose is to see how the inclusion of the UK into this group affects this common cycle, and how the idiosyncratic parts of the cycles interact with it. The results suggest that adding the UK to the Euro group does not lead to a greater heterogeneity of this group. In addition, we find that the UK output cycle is more correlated with the US cycle but is increasingly synchronized with the other European cycles.

The last part is somewhat different. This paper is built upon a monetary, new Keynesian model developed by Laubach & Williams (2001). This model tries to estimate the natural real interest rate for the Euro Area. The new Keynesian nature of the model implies that 'business cycles'[6] determine directly the level of inflation, through a Phillips curve-like relation. At the same time, there is a mutual influence between the 'business cycle' and the natural rate of interest. As such, this paper provides an example of how business cycles can interact with other macroeconomic variables. Since Kalman filtering techniques are used, the definition of cycles correspond more to the one used in the third part.

[6]In reality, the model uses output gaps instead of business cycles. Although the two concepts are often used indifferently in the literature, they are not measured in the same way. The output gap is simply the difference between the series and its long run trend, or *potential output*. Therefore, it represents the elements of the series that are unexplained by long run movements. It incorporates high-frequency elements that are normally not part of business cycles. The reason why the two concepts are often mixed up is that output gaps can be seen as adequate proxies for a certain type of business cycles, that corresponding to the trend/cycles decomposition.

The difference between the two elements is also conceptual in a sense. Business cycles are modelized explicitly –see for example the third paper– while output gaps are only defined indirectly through a third element, i.e. the trend. At the same time, the idea of business cycles is more general than that of output gap, since it groups together different definitions, as we have seen above.

Chapter 1

Has the Similarity of Business Cycles in Europe Increased with the Monetary Integration Process? A Use of Classical Business Cycles

Abstract

We investigate to what extent the business cycles in Europe have become more synchronised since the sixties, using the classical business cycles framework. Different Bry & Boschan-like procedures for dating the turning points are compared. It is found that our univariate procedure performs as well as others, by comparing turning points dates with those found by the more sophisticated procedures of the NBER or the ECRI research center. Another point is that there are great differences in the dates found from one procedure to the other. Concerning cycles across countries, we find that they have become more idiosyncratic through time, but also that it is less obvious for the Euro countries. The main conclusion is the existence of a core group within the Euro area with more strongly linked cycles.[1]

[1]This chapter is a modified version of Garnier (2003)

1.1 Introduction

This paper is based on the classical business cycle framework. Its aim is to see whether the creation of the European Monetary System has been correlated with a greater similarity of the business cycles across Europe. In other words, the question I will try to answer is: has the nature of business cycles been modified from the *pre* to the *post* EMS period, and have those cycles become more similar? This is an important question since the homogeneity of the cycles may be seen as one of the requirements for the Euro to be an adequate instrument. This work can be directly related to the more general debate on Optimal Currency Areas (OCA) and the Lucas critique. If it is found that business cycles in the European Union are homogenous over the whole period, this would tend to show that the EU is *intrinsically* an OCA. Inversely, we might find that the business cycles exhibit no homogeneity at all, and consequently that the EU is not an optimal monetary zone. If it is found that the cycles have become more similar through time, this would tend to suggest that the monetary integration process has had some influence on the homogenisation of the business cycles. The implication would be that this process in itself increases the probability of being an OCA (Frankel & Rose, 1998).

The classical business cycle approach deals with cycles in levels. We are required to find the turning points (henceforth TPs) first. We will denote the peaks by P and the troughs by T. For the US, the turning points identified by the NBER are often regarded as the 'official' turning points. A similar dating procedure is used by the ECRI[2] for several other countries. The EUROCOIN indicator published by the CEPR provides a dating for the whole Euro area[3]. The Bry & Boschan algorithm (BB) is a practical tool for replicating such dating procedures. However, it is a univariate tool, which implies that it cannot be substituted to the more sophisticated approaches of the NBER or the ECRI. We will see below that this feature can be problematic. Its basic rule is that a point in

[2]Economic Cycles Research Institute (www.businesscycles.com). This is a private organisation working on the analysis and the forecast of business cycles. To my knowledge, this is the only publicly available alternative to the NBER for dating (classical) business cycles.

[3]See www.cepr.org. The EUROCOIN indicator is based on an econometric model that looks after common components within a large set of stationarized time series. The output is a 'common business cycle' within which turning points are selected. Since the cycle is stationary, this approach refers more to 'growth' than 'classical' cycles. Note that the OECD publishes TP dates, but based on the *Phase Average Trend* method, which is also closer to the growth than the classical cycles framework.

t is a turning point if it is the highest/lowest point within a period of $t \pm n$. We will use here a modified version of the procedure used in the article of Artis, Kontolemis & Osborn (1997, henceforth AKO), inspired by the BB algorithm. All these procedures are univariate applications.

Once the turning points have been found, it is possible to compute the phases of the series, i.e. the expansions (T-P) and the recessions (P-T), and to start studying the evolution of the cycles. The approach used is non-parametric and based on descriptive methods. At first, we will look at the shapes of the cycles and see if they have become more similar through time. Subsequently, the *timing* of the cycles will be considered. The idea is to see how the expansion/recession phases are coordinated across countries.

The series used for this study is the seasonally adjusted index of industrial production, provided by the OECD. The data set comprises 18 countries and the sample starts in January 1962 and ends in January 2001. The panel is composed of the 12 Euro countries minus Ireland[4], three countries that belong to the EU but not to the Euro (Denmark, Sweden and the UK), two European countries outside the EU (Norway and Switzerland) and two 'external' countries (Japan and the US). This division of the panel into different areas should make it easier to evaluate the influence of the monetary integration. One might believe that if this influence exists, it should be more important for the Euro group than for the external countries. We have taken this classification to ease the presentation. However, one should be careful in interpreting the results. Indeed, half of the Euro countries did not belong to the EEC at the beginning of the sample and most of them only entered in the second period (Greece in 1979, Spain and Portugal in 1986, Austria and Finland in 1995). The positive aspect of this is that if the monetary integration has some influence on business cycles, we should find weaker results for the 'latecomers'. In the following, we will make a distinction between countries that belong to the EEC from the beginning of the sample (the former Federal Republic of Germany, France, Italy, Belgium, the Netherlands and Luxembourg) and the others.

To assess the evolution of the cycles, the sample will be divided into two sub-samples, before and after March 1979. This date corresponds to the creation of the European

[4]That is: Austria, Belgium, Finland, France, Germany, Greece, Italy, Luxembourg, the Netherlands, Portugal and Spain. Concerning Ireland, the shape of its industrial production series makes the finding of turning points difficult. It would be more useful in this case to study the growth cycle.

Monetary System (EMS) and of the Exchange Rate Mechanism (ERM), which is the first[5] real attempt to create an explicit monetary system at the European level. This date can be seen as the starting point of the monetary integration.

The classical business cycle framework, initiated by the empirical work of Burns & Mitchell (1946), has recently been the subject of a revival of interest following the articles of Harding and Pagan (2000b, 2002, 2003) or Artis et al. (1997). See Artis et al. (2003) for an extension of the methodology. The procedure adopted here is based on AKO, with the period of analysis updated to January 2001. As Harding & Pagan point out, the advantage of using the classical cycles approach is the freedom from arbitrary assumptions about the trend. Indeed, one of the problems of 'detrending' (or 'filtering') techniques is that their results differ from one another (Canova, 1998). In particular, it has been argued that ad-hoc filters could create spurious cycles[6]. Here, the method does not remove any trend as it deals with cycles in levels. The next part is dedicated to the exposition of the procedure used to delimit those classical cycles. The procedure is also compared to other dating methods. In the third part, we use the type of plots presented by Burns and Mitchell (1946). These plots display the average cycles of the series considered. A more recent utilisation of this technique can be found for example in King & Plosser (1994) and Simkins (1994). The fourth part investigates the evolution of the *timing* of the cycles, and the last part concludes.

1.2 Finding turning points

1.2.1 Description of the procedure

The procedure[7] used here aims at replicating the BB-like procedures and in particular the one by AKO. We will try to show in the following part that our simplified version of the latter might be as efficient as the other procedures in capturing turning points in an industrial production index series.

[5]In fact, the 'European Snake', created in 1972, was already an attempt to create a certain homogeneity among the currencies of the European countries, but it had been created in the context of the Bretton-Woods system.

[6]King & Rebelo (1993), Osborn (1995) or Harvey & Jaeger (1993) provide such results for the Hodrick-Prescott filter. A good overview of the problem can be found in Guay & St Amand (1997).

[7]The codes, written in GAUSS, are available upon request.

There are no major differences between the BB, AKO procedures or the one used here. Turning points are essentially based on the selection of local peaks and troughs, which are the highest/lowest points within periods of several months before and after them. Peaks and troughs are also required to alternate. Among those local turning points, the final dates are selected by various procedures. See appendix A.1.1 for more details. Artis et al. (2003) propose an interesting development of this approach. Their procedure is based upon transition probabilities obtained from Markov chains. Despite this conceptual difference, it finds the same turning points as Harding & Pagan (2001) for the Euro area.

The algorithm proceeds in four main steps (see the appendix for details). The first one determines the outliers, i.e. the points x_t such that : $(x_t - x_{t-1}) \geq 3.5\sigma_x$, where σ_x is the standard deviation of x_t. These outliers are replaced by the average of the two adjacent observations[8]. It shall be noticed that doing so might be problematic. Indeed, this comes down to transforming arbitrarily the data, which may bias resulting turning points. However, leaving outliers might bias the results even more, so that we follow the literature on this point. Step 2 finds the turning points in the series smoothed by a 12-months moving average. The smoothing allows to get rid of idiosyncratic fluctuations that could modify the results. Step 3 uses the raw series to find the turning points. Then, short cycles (less than 15 months from peak to peak or trough to trough) are identified and eliminated, by keeping the highest (lowest) of the two peaks (troughs). Finally, each phase (P-T of T-P) is required to have an amplitude of at least one standard error of the series considered. When the procedure meets a phase of low amplitude, it eliminates its last turning point and keeps the first. I have followed Watson (1994) and Harding & Pagan's programs on this point. The last step compares the dates found in step 2 (smoothed series) and step 3 (raw series) and states the final set of turning points.

There are several differences between the AKO procedure and the one used here. The main one is that the identification of flat segments (in step 3 of AKO) has been suppressed here. There are two reasons for this. First, it is not really justified by the authors. Second, I conjecture that this is not necessary, because of the requirement, in step 3 of the procedure used here, that each phase should have an amplitude of one standard error. This should produce the same result. Another difference is that the enforcement of alternation

[8] Mark Watson uses the value given by the Spencer curve for that observation. The difference between the two corrected values should be marginal.

has been placed at the end (step 4a), whereas it was used twice in the AKO procedure. Note that I had followed the AKO procedure on this point at the beginning, but the results were exactly the same as the ones presented here.

1.2.2 Results

Summary of the results for the dating of turning points (TPs)

Number of TPs found by ECRI*	86
Number of TPs found here**	111
Number of TPs found by BBW**	153
Proportion of ECRI dates captured by BBJG***	48.80%
Proportion of ECRI dates captured by BBW***	60.40%
Proportion of BBJG TPs well-identified***	37.90%
Proportion of BBW TPs well-identified***	34.50%
Total number of TPs in common between BBJG and BBW	167
Proportion of TPs of BBJG found by BBW	62.50%
Proportion of TPs of BBW found by BBJG	88.40%

Sample period Jan.1962 - Jan 2001
nb: the TPs found by AKO are not reported because their sample is shorter
BBJG : procedure used here *BBW* : BB proc written by Watson (1994)
* for 10 countries
** same countries as the ECRI
*** a date is 'well identified' if it is not distant by more than one term from the ECRI or the NBER ones

Table 1.1: Ability of univariate procedures to find adequate turning points

Table 1.1 shows summary results for the turning points[9]. The dates found by different procedures are compared. We compare the dates found here ('TPs found by BBJG') with those of Watson (1994) who replicates the BB procedures. The turning points published by the ECRI are taken as a benchmark. We take into account the fact that the algorithms do not find the same number of peaks and troughs. Univariate BB applications identify more dates than the ECRI, and Watson's procedure more than ours (rows 1 to 3 of the table). The quality of a procedure can be evaluated on two aspects: First, the proportion of the total number of 'good' dates[10] that are captured by the procedure (rows 4-5 of the

[9]Complete results are displayed in the appendix (table A.1 and following tables). We make a comparison with the dates found by the BB procedure of Watson (1994) and –when available– with those published by the ECRI. As the ECRI uses the same approach as the NBER, we can use it as a good benchmark to assess the other dating procedures. When there is a correspondence, the dates found by AKO are shown as well.

[10]As a criterion, we take dates that are not distant by more than one term from the ECRI ones.

table). This would be given by $N_{\mathrm{BB}\cap\mathrm{ECRI}}/N_{ECRI}$, where N_p is the total number of turning points found by the procedure p and $BB \cap ECRI$ is the set of turning points in common between BB and $ECRI$. Second, the proportion of dates found by the procedure that are 'good' ones: $N_{BB\cap ECRI}/N_{BB}$ (rows 6-7). Since the first measure tends to increase with the number of dates identified, we prefer the second one. The Watson's BB procedure outperforms our procedure for the first measure, but not for the second. See paragraph below.

The turning point dates differ from those found by AKO, although the source of the data is the same (OECD). There are great differences from one country to another. For some of them, the results are similar (Germany, Italy, Luxembourg, the Netherlands, the UK, Japan, the US[11]), and for others they are quite different (Spain, Belgium). The origin of these differences might be twofold. A) There can be differences in the dataset. I have used my procedure with the dataset of AKO, but the turning points were still divergent. The fact that the sample periods are not the same for the two datasets[12] could explain some of the differences for the early nineties. That is, the procedure keeps the highest (lowest) of two consecutive peaks (or troughs), because of the alternation (P-T-P...) requirement. For example, if the last turning point is a trough, we might find another trough immediately after if the sample was extended. If one states that it is lower than the last 'in-sample' one, then it would be selected at the expense of the previous one. Data revision has also occurred for some countries. In order to verify this, I have plotted for each country the series from the two datasets[13]. B) The differences in the dates can be due to differences in the procedures themselves (see the paragraph above). I have taken the same dataset as AKO with the two procedures. Some differences remain. Nevertheless, similar dates are found for most countries.

The table above reveals that most of the TP dates found here are also identified by BBW but the latter finds more dates than our procedure. Either our procedure does not capture enough dates or BBW captures too many of them. For some countries (e.g. Switzerland) the dates are quite similar between our procedure and the BBW, but com-

[11]One date is different from the AKO. I have found a peak in May 1979 and AKO in March 1980. Apparently, some data revision has occured. May 1979 is actually higher than March 1980 in data used here (which was not the case in the dataset used by AKO).

[12]1961:1 - 1993:12 for Artis et al. (1997), and 1962:1-2004:12 here

[13]These graphs are not reported here but are available on request.

pletely different from the dates of the ECRI. For a country like the UK, BB-like procedures identify almost all the dates of the ECRI, but they also find more dates. Overall, the BB procedure of Watson finds 1.78 times more turning points and the one used here 1.29 times more, which suggests that they both overidentify TPs.

We see that only half –50 to 60% – of the 'true' turning points are captured by the procedures (rows 4-5). Besides, about one third of the TPs found are 'true' ones (rows 6-7). This is a poor result at first sight, but one has to keep in mind that the ECRI (and the NBER) have a global view of the economy[14], whereas the BB procedures used here are only univariate.[15] This points out that one has to be careful when interpreting the results. The fact that algorithms of the BB-type extensively use the rule of the highest/lowest points (e.g. if there is a choice between two peaks, the highest one will be selected), implies that the turning points do not necessarily coincide. In other words, the highest/lowest of two points may not necessarily be the same for the industrial production and the GDP series, even if the occurrence in time of a particular event is exactly the same. Making the assumption that the algorithms are not ill-defined and that it is not incorrect to apply BB-type procedures onto industrial production series, we can say that our procedure does a slightly better job than BBW. It is true that BBW captures more 'true' TPs than our procedure, but at the same time the total number of TPs identified is much higher. At the limit, a procedure that would capture *every* date of the sample would also capture all the true dates. It is more interesting to look at the proportion of true TPs amongst the ones identified by the procedure. This proportion is slightly higher for our procedure than for BBW (37.9% against 34.5%) . For this reason, we will prefer our procedure for the remaining parts of this paper.

1.3 Comparisons of cycles based on their shapes

We use here the type of plots used by Burns and Mitchell (1946) and also by King & Plosser (1994) and Simkins (1994). The idea is to make a representation of the typical classical

[14] The ECRI uses several macro series –essentially output, income, employment and sales. For each of them, the Bry and Boschan procedure is computed and the final turning points are chosen on "*the basis of the best consensus*" among the different series.

[15] Artis (2002) uses the procedure of Artis et al. (2003) on a *monthly* GDP series for the UK and finds remarkably similar turning points as the ECRI.

cycle of a series. For one particular country, each phase (delimited by two turning points. P-T or T-P) is divided into four sub-periods and the average growth rate and duration of these sub-periods are taken. The average expansion and recession phases are finally put together in order to make the graph, which has the form Trough-Peak-Trough.

For the first and last sub-periods of each phase, we take three months after the first turning points and three month before the second. The time in between is divided in two equal parts.

To avoid the bias that could result from the idiosyncratic movements and from the small number of elements included in the average, each series has been smoothed by a moving-average. A centred MA(7) has been arbitrarily taken. This choice is motivated by two reasons. First, each point must not capture too much information from the past and future observations. That is, the elements of one sub-period should not be substantially influenced by the elements of the adjacent sub-periods. Second, excessive idiosyncratic movements must be smoothed sufficiently. As a centred MA(7) takes the information one term before and one term after t, it seemed a good compromise between those two points.

The results of these graphs are shown below and in the appendix. Note that some countries are absent. This was the case when not enough turning points were found in the period considered –that is, less than three. The sample has been divided in two in March 1979 (date of the creation of the ERM). If the monetary integration has had an influence in the second period, one should find a greater homogeneity in the shapes of the cycles within the group of countries that have moved towards this integration. Another possibility is that if the Euro area is intrinsically an OCA, the shapes of the cycles should be similar before and after 1979. Under this hypothesis, there should be differences between the shapes of the countries belonging to the OCA and those of the other countries. Of course, it is also possible that no clear pattern appears from the plots.

Overall, one has the impression that the shapes are quite heterogenous. However, if we look more specifically at the Euro group, a certain degree of homogeneity can be oberved in the second period, while the first shows no clear pattern (figures 1.1 and 1.2). In the second period, it seems that the cycles are more similar in duration and growth rates, with the exception of Portugal and Luxembourg. Conversely, the cycles shape of 'non-Euro' countries are quite dissimilar. Note that plots of the whole EU are produced in the appendix. They show that adding 'EU-non-Euro' countries cycles to the Euro group increases the heterogeneity of the group, for both periods.

Figure A.9 in the appendix provides the same plot for all the Euro countries as well as the separate plots of each cycle. If one looks at expansions for the second period, three groups appear. The first, for which the expansion period is slightly slower at the beginning than at the end of the phase, is composed of Austria, Greece and Italy. In the second group, we find the opposite pattern, with a stronger growth first. This is the case of Belgium, Luxembourg, the Netherlands, Spain and to a lesser extent, Portugal. It is interesting to note that a strong growth that slows at the end is also a characteristic of the US expansion phase often found in the literature. France and Germany exhibit an almost linear expansion period, so that they lie somehow between the two groups. Note that their shapes are very similar except that recessions are more severe in the former.

Business cycles lengths are also interesting. They are more homogenous in the second period. Once again, two groups appear. One group has cycles of length comprised between 60 and 75 months: Belgium, France, Germany, Italy, Luxembourg and the Netherlands. The second one, composed of Austria, Greece, Portugal and Spain shows longer cycles –more than 100 months.

Recessions do not provide much information for no pattern can be seen: their shapes differ too greatly from one country to the other. At the same time, their duration is about the same for most of the countries of the panel, around 20 months (between 15 and 40 months).

One should be aware of the limitations of this approach. The averages are calculated with very few elements and some strange behaviours can be observed[16]. At best, these graphs can reinforce a prior intuition, but they are not sufficient.

It seems that a Euro pattern has appeared among a core group. However, such conclusions are tempered by the fact that only a small number of phases enter into the calculations. We now have to go beyond those first impressions. We will see whether they are confirmed by more precise measurement of the *timing* of the cycles.

[16]Belgium in the first period for example. This is due to the fact that a phase is strictly defined as a period between two TPs, such that the period before the first turning point cannot be taken into account. For Belgium, the period of stagnation –observed on the stylized plot– is in fact preceded by a period of growth of six years, such that in reality, the pre-EMS period for this country is characterized by an increase of industrial output and not by stagnation.

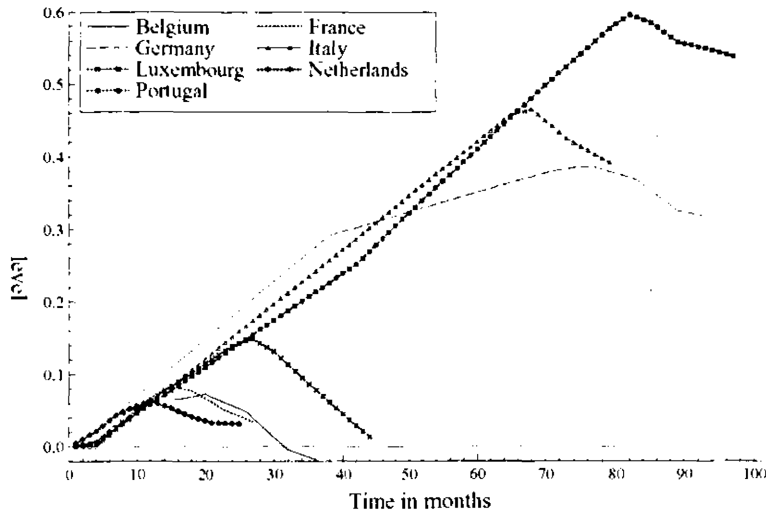

Figure 1.1: Typical cycles - 'Euro' group - First period

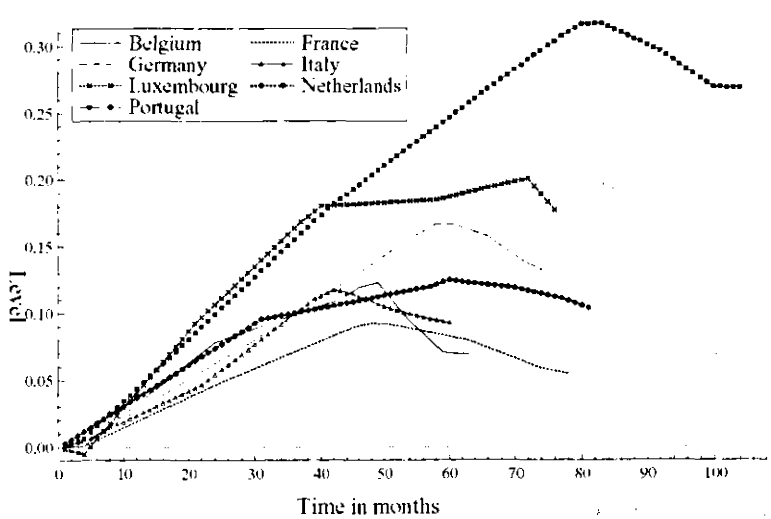

Figure 1.2: Typical cycles - 'Euro' group (same countries as first period) - Second period

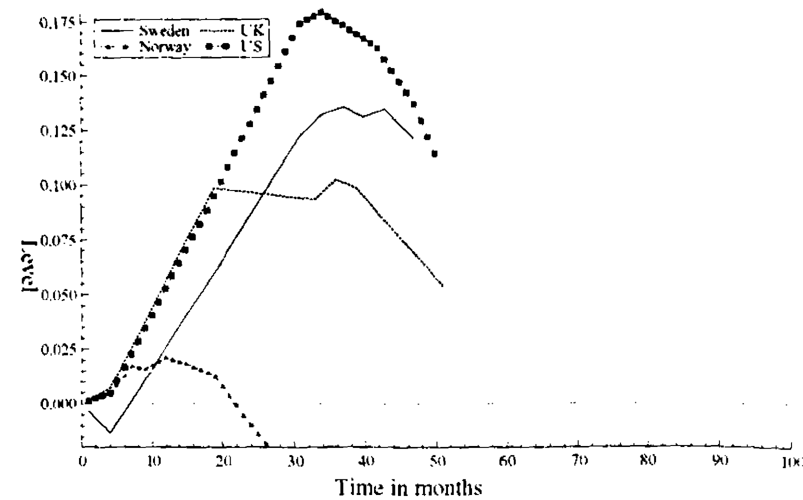

Figure 1.3: Typical cycles - 'External' group - First period

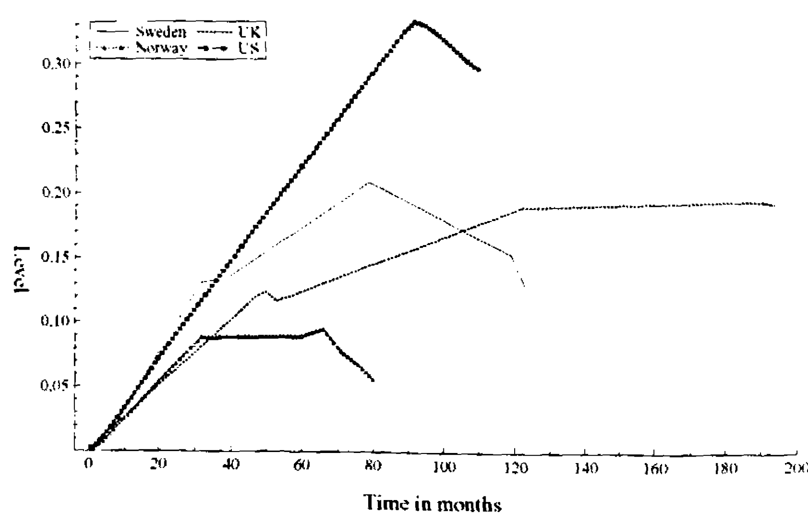

Figure 1.4: Typical cycles - 'External' group (same countries as first period) - Second period

1.4 Comparisons of cycles based on time synchronization

We study in this section the comovements of the cycles from one country to another. We ask the question of whether the European countries have become more synchronised or not. To see this, we will take two countries as references, the US and Germany[17], in the same spirit as Artis & Zhang (1996). The assumption is made that this latter country leads the European economy and that if a European cycle exists, it should be affiliated to the German cycle. The idea is to check whether European integration has led the European countries to a greater synchronisation with Germany than with the US. If we find that the series have become increasingly more dependent from the former than from the latter, we could suspect a relation between the implementation of the EEC / EU and the synchronisation of the cycle.

Two methods are computed to evaluate the comovements: Pearson's coefficient and Harding & Pagan's concordance ratio. Each one considers two binary series (expansions / recessions) at a time and the number of periods where these series are in the same phase. Each one evaluates also how much the series depart from the null of perfect independence. Two steps are needed for the second one, but it has the advantage of always being computable, which is not the case of the first method.

Harding & Pagan's index is specifically designed for business cycles and has been used several times in the recent literature (e.g. Artis et al., 2003). In a sense, it measures the same thing as the Pearson's coefficient and there is some redundancy in using both of them. However, there are sensible differences in the results. Indeed, they use different formulas even though both of them are built upon the same principle –comparing a particular state of the nature with the one that would be expected under perfect independence. The main difference between the two approaches is that the Pearson's coefficient takes all the four possible states while Harding & Pagan's concordance index only looks at cases where the series are in the same phase.

In order to capture the information from the periods *before* the first and *after* the last turning points, a procedure has been added here: if the first point of the sample is higher (/lower) than the first turning point and if this one is a trough (/peak), the procedure creates an 'artificial' peak (/trough) at the first observation. The equivalent is done at the

[17] Calculations involving all the pairs of countries have also been computed. They have not been reproduced here for more clarity, but are available upon request.

end of the sample. This procedure has been essentially designed to take into consideration the two periods of expansions, in the early 60s and in the 90s, situated at the extremities of the sample. They would be eliminated otherwise and the results would be biased. Note that this procedure could obviously not be applied in the study of shapes above.

Let s_{it}, $t = 1, ..., T$ a dummy equal to one when the industrial production series of country i is in expansion and zero otherwise. Define also $n_i^1 = \sum_{t=1}^{T} s_{it}$ the number of periods where country i is in expansion, and $n_i^0 = \sum_{t=1}^{T}(1-s_{it})$ the equivalent for recessions. Let n_{ij}^{rs} the number of periods where i is in phase r and j in phase s. Finally,

$$n_{ij} = n_{ij}^{11} + n_{ij}^{00} = \sum_{t=1}^{T}[s_{it}s_{jt} + (1-s_{it})(1-s_{jt})]$$

is the number of periods where countries i and j are in the same phase. We see that $n_{ij}^{r\cdot} = n_i^r$.

		Country j		
		Expansion	Recession	Subtotal
	Expansion	n_{ij}^{11}	n_{ij}^{10}	$n_{ij}^{1\cdot}$
Country i	Recession	n_{ij}^{01}	n_{ij}^{00}	$n_{ij}^{0\cdot}$
	Subtotal	$n_{ij}^{\cdot 1}$	$n_{ij}^{\cdot 0}$	T

1.4.1 Pearson's coefficient

- Methodology :

This coefficient is based upon the chi-square statistic:

$$\chi_{ij}^2 = \sum_{r=0}^{1}\sum_{s=0}^{1}\frac{[n_{ij}^{rs} - n_{ij}^{r\cdot}n_{ij}^{\cdot s}/T]^2}{n_{ij}^{r\cdot}n_{ij}^{\cdot s}/T} \tag{1.1}$$

In a sense, this expression measures the difference between the actual and the expected number of periods where country i is in phase r and country j in phase s. The expected value of n_{ij}^{rs} is the one occuring when the two series are perfectly independent. It is given by $n_{ij}^{r\cdot}n_{ij}^{\cdot s}/T$.[18] Under the null of independence between the two series, eq.(1.1) follows a

[18] This expression can be interpreted as the total number of periods where i is in state r, $n_{ij}^{r\cdot}$, times the probability for j to be in state s, $n_{ij}^{\cdot s}/T$. In other words, the pearson coefficient does not assign the same probability to each state. For this reason, the Pearson's coefficient must be able to take into account trending series, for which the probability of being in expansion is higher than the probability of being in recession.

chi-square distribution. If the actual value n_{ij}^{rs} is well above the expected one, we might suspect the existence of a statistical link between the series i and j. That is, the occurence of one state (here expansion or recession) for one country would be associated to another particular state for the other country more often than in the case of independence. The size of the Pearson's coefficient depends upon the strengh of the relation between the two variables. A problem is that it depends on the sample size as well. Pearson's contingency coefficient, which corrects for the sample size bias should therefore be used.

$$CC_{ij} = \sqrt{\frac{\chi_{ij}^2}{T + \chi_{ij}^2}} \tag{1.2}$$

An additional problem is that for finite dimensions, the contingency coefficient is bounded above and is biased from its true value. This limit is proportional to the dimension of the table. Here, only two variables are considered. Therefore, the bias might be quite high. As the limit of the coefficient for such a dimension is $\sqrt{1/2}$, the corrected coefficient is given by:

$$CCC_{ij} = \frac{CC_{ij}}{\sqrt{1/2}}$$

We can observe that some values are missing in the tables of Pearson's coefficient (see the appendix for detailed results). This happens with small sub-samples when all the possible cases are not present. For example, if a country did not experience any recession during the period studied, $n_{ij}^{r.} = n_i^r$ must be null too (e.g. the expansion period in the US after March 1991). By eq.(1.1), this is impossible.

- Some results :

Table 1.2 shows the corrected Pearson's coefficients for all the countries with Germany and the US[19]. Figures 1.5 and 1.6 show the same results.

Note that the coefficient between Germany and the US has declined between the two periods. This might constitute a first argument towards the autonomy of the European cycle. The coefficient goes from 0.54 (or 0.46, depending on which country is taken as independent) in the first period, to 0.27 (0.34) in the second.

, In the first period, most of the countries are grouped around the 45° line, whereas there is a strong movement towards the German cycle in the second. The case of France

[19]Note that the coefficients of Germany (or US) with itself is 0.99 in period 1 because of roundings

Pearson's coefficient

	Period 1		Period 2	
	Germany	*US*	*Germany*	*US*
Au	0.69	0.64	0.89	0.44
Be	0.92	0.71	0.48	0.12
Fi	0.67	0.61	0.12	0.41
Fr	0.59	0.61	0.86	0.42
Ge	0.99	0.54	1.00	0.30
Gr	0.43	0.47	0.83	0.45
It	0.39	0.44	0.36	0.54
Lux	0.57	0.34	0.50	0.15
Net	0.64	0.58	0.80	0.36
Po	0.78	0.75	0.41	0.11
Sp	0.52	0.56	0.28	0.25
De	0.00	0.00	0.31	0.36
Swe	0.37	0.50	0.13	0.02
UK	0.58	0.21	0.34	0.12
No	0.59	0.51	0.51	0.60
Swi	0.57	0.27	0.23	0.29
Ja	0.69	0.73	0.39	0.05
US	0.47	0.99	0.35	1.00

Table 1.2: Comovements of countries *vis-à-vis* Germany and the US

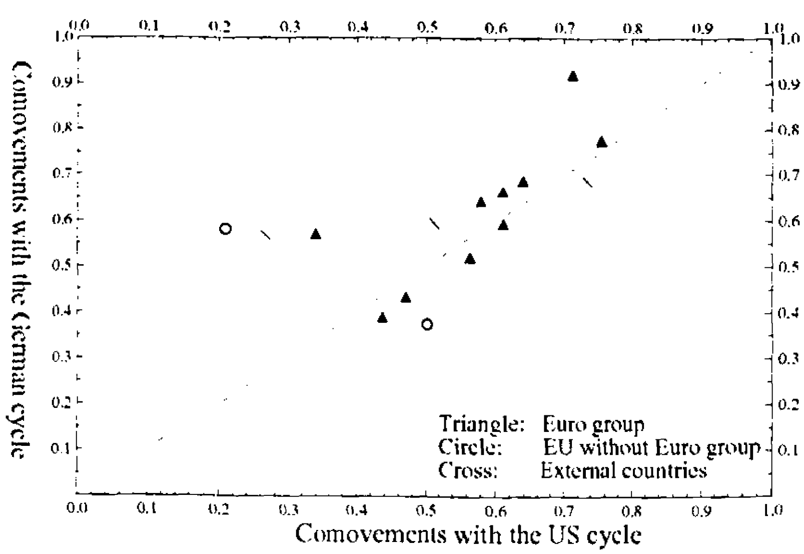

Figure 1.5: Pearson's corrected coefficient - First period (1962 - 1979)

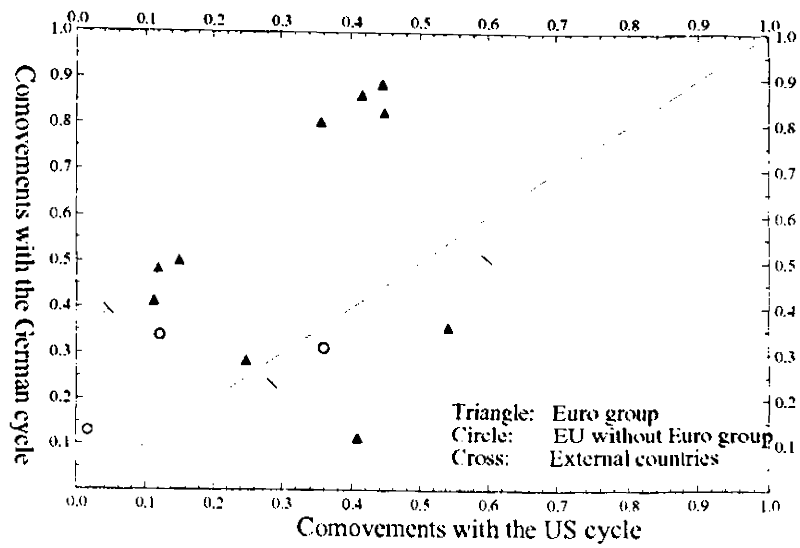

Figure 1.6: Pearson's corrected coefficient - Second period (1979 - 2001)

is quite representative. In the first period, its coefficient is higher with the US than with Germany, but the difference is small (0.59 against 0.61). In the second, it is much more correlated with the German cycle (0.86 against 0.42). The fact that two countries of the Euro group are more synchronous to the US cycle in the second period suggests the existence of two groups within the Euro area. Those countries are Finland and Italy. There is a possibility that countries which are not under the influence of Germany can be 'caught' by the American cycle.

If we look at the Euro countries that are *above* the line –i.e. Euro countries minus Finland and Italy–, we see that the level of the coefficients *vis-à-vis* the German cycle is stable between the two periods (0.68 to 0.67 on average). At the same time, the coefficient with the US decreases in the second period (0.58 to 0.29). Therefore, it seems that the links within a 'core' group of Euro countries have remained stable at a high level whereas there is greater independence towards the US cycle.

Concerning external countries, the UK is more correlated with Germany in both periods, unlike Sweden, which has a higher coefficient with the US (the Denmark coefficient cannot be calculated in the first period because only one turning point was captured). Note the place of Japan, which has been taken as a control country. Its coefficient against

Germany and the US is the same in the first period. In the second period it is more correlated with the German than the US cycle. This result it quite surprising as the Japanese economy is rather closed, and it exchanges more with the US than with Western Europe.

The fact that the Pearson's coefficient cannot be computed in some cases limits its impact. We will see next the Harding & Pagan's concordance that has the advantage of always being computable.

1.4.2 Harding & Pagan's concordance index

- Methodology :

The second index is built specifically for business cycles analysis (Harding & Pagan, 2002). It compares the number of periods where the two series are in the same phase with the expected number of periods under the null of independence between these two series. That is, it takes into account n_{ij}^{11} and n_{ij}^{00} while Pearson's coefficient takes the four states, considering also periods where the two series are not in the same phase. The index proceeds in two steps. The first one calculates the *concordance index* strictly speaking. The second step makes a ratio with the expected index, computed under the null of independence. The index is given by:

$$I_{ij} = \frac{n_{ij}}{T} = \frac{1}{T}\sum_{t=1}^{T}[s_{it}s_{jt} + (1 - s_{it})(1 - s_{jt})] \tag{1.3}$$

The index shows how the specific variable i behaves in relation to the reference series, j, here Germany or the US. If the index is one, country i is exactly pro-cyclical with respect to j. Conversely, if it is null, the index indicates a counter-cyclical series.

If the two series are statistically perfectly independent, the expected index is equal to the probability that the series happen to be in the same phase at a given time t :

$$E[I_{ij,t}] = E[s_{it}]E[s_{jt}] + (1 - E[s_{it}])(1 - E[s_{jt}]) \tag{1.4}$$

Each expectation can be measured by the number of time units where the state occurs, divided by T. It is then easy to compare the concordance and the expected indexes. If the former is higher than the latter, it can be said that there is a link between the cycles, because the number of periods where the series are in the same phases is higher than if the series were totally independent. Conversely, if the ratio between the two is less than one, we can suppose that there is a counter-cyclical relation between the two series. Of course, we cannot say anything about the level of this ratio, and it would be better to derive some test to see if the ratio is significantly different from one. This is the topic of the next section.

- Some results :

As before, we only give here the coefficient between the US, Germany[20] and the other countries. Complete results are given in the appendix.

Harding & Pagan's ratio

	Period 1		*Period 2*	
	Germany	*US*	*Germany*	*US*
Au	1.18	1.13	1.58	1.21
Be	1.48	1.37	1.33	1.07
Fi	1.17	1.12	1.03	1.11
Fr	1.13	1.12	1.53	1.21
Ge	1.54	1.21	1.78	1.15
Gr	1.06	1.06	1.57	1.23
It	1.10	1.11	1.21	1.33
Lux	1.29	1.15	1.30	0.92
Net	1.16	1.10	1.52	1.18
Po	1.25	1.20	1.18	0.95
Sp	1.13	1.14	1.09	1.06
De	0.00	0.00	1.13	0.87
Swe	1.12	1.15	0.93	0.99
UK	1.25	1.07	1.24	0.92
No	1.13	1.08	1.28	1.35
Swi	1.15	1.05	0.86	0.83
Ja	1.18	1.19	1.23	0.98
US	1.18	1.47	1.18	1.65

Table 1.3: Comovements of countries *vis-à-vis* Germany and the US

The ratio between the US and Germany goes from 1.2063 in the first period to 1.1521 in the second. As for the Pearson index, the German economy is slightly more independent from the US in the second period.

Alike the previous part, we see on figures 1.7 and 1.8 the general movement of the Euro countries towards Germany in the second period. Note that the scale of these plots should not be compared to the previous ones. As before, two Euro countries are more

[20] The coefficients for these countries with themselves are not integers since the Harding & Pagan ratio is not bounded above. Note however that these are the highest numbers of each column

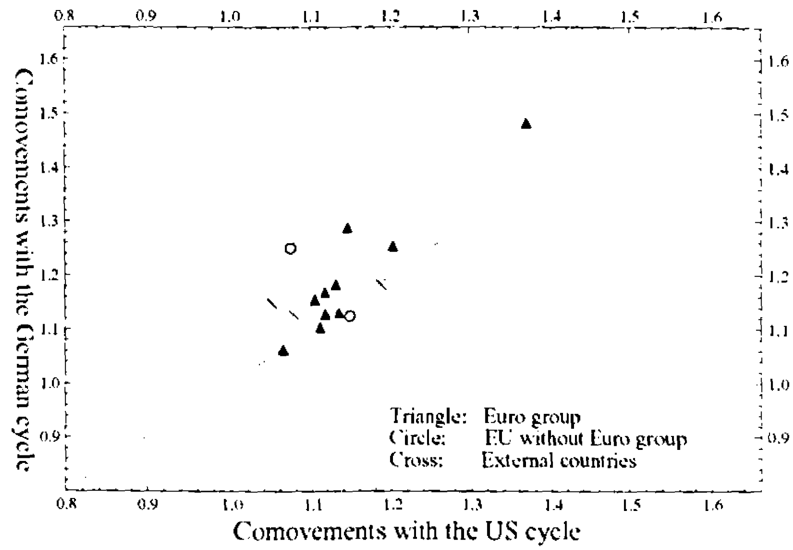

Figure 1.7: Harding & Pagan concordance ratio - First period (1962 - 1979)

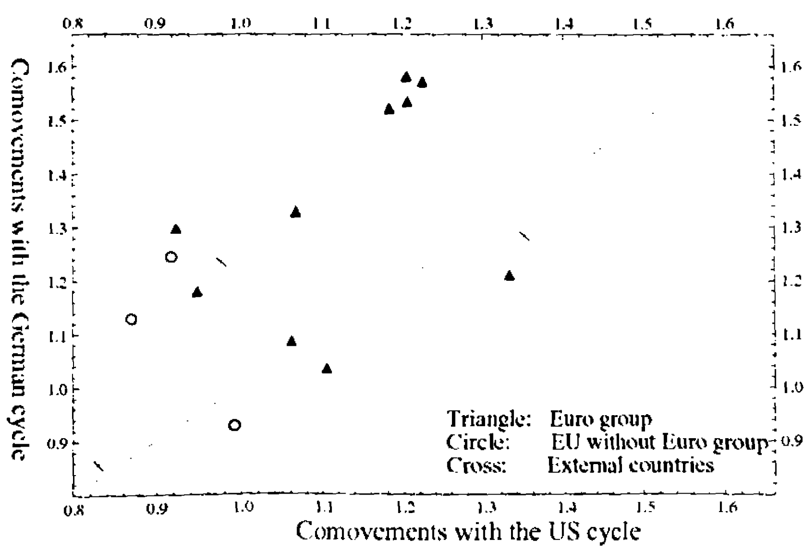

Figure 1.8: Harding & Pagan concordance ratio - Second period (1979 - 2001)

correlated with the US than with Germany: Finland and Italy. Note once again the place of Japan, which is closer to the German than the US cycle.

If we look at the ratios of the Euro countries –excluding again Finland and Italy–, the average comovements with Germany is 1.25 whereas it is 1.16 for the US in the first period. In the second period, the ratio increases to 1.43 for the former and decreases to 1.11 for the latter. This suggests that the dependence with the US cycle has decreased, whereas it has increased *vis-à-vis* the German cycle. If one looks at the average comovements for the group of European countries that do not belong to the monetary system[21], the figure with respect to Germany decreases (1.19 to 1.09) but less than with respect to the US (1.11 to 0.96). Therefore, it seems that the European cycles have acquired greater independence towards the US.

1.4.3 Attempt to test the concordance (Harding & Pagan 2000a)

Testing the independence between the phases of two series

The measures of comovement we have seen above have the disadvantage of not being meaningful alone. Using benchmarks –as Germany and the US– is needed in order to do comparisons. What would be helpful at this stage would be a test that would tell us whether or not there is actually a relation between phases. Consequently, we will try in this section to implement the test suggested by Harding & Pagan (2000a, p.11). The idea is to take a binary variable representing the phases of a series (expansion/recession) and to regress it on another variable of the same type. The null hypothesis is that there is no statistical relation between them. Under the null, the corresponding coefficient should be zero.

To see why this test is consistent with the approach of Harding & Pagan, consider the following. Eq.(1.3) can be rewritten as :

$$\hat{I} = \frac{1}{T}\left[2\sum_{t=1}^{T} s_{it}s_{jt} + T - \sum_{t=1}^{T} s_{it} - \sum_{t=1}^{T} s_{jt}\right] \tag{1.5}$$

Similarly, equation 1.4 becomes :

[21] i.e. Sweden and the UK only, since the ratio for Denmark is equal to zero in the first period

$$
\begin{aligned}
E[I_{jk}] &= 2E[s_{it}]E[s_{jt}] + 1 - E[s_{it}] - E[s_{jt}] \\
&= 2\hat{\mu}_i\hat{\mu}_j + 1 - \hat{\mu}_i - \hat{\mu}_j \qquad (1.6)
\end{aligned}
$$

for the sample considered.

As we have seen above, (1.5) and (1.6) are equal if the two series are perfectly independent. So the *mean corrected* concordance equals zero under the null of independence, and is :

$$
\begin{aligned}
\hat{I}_{mc} &= \frac{1}{T}\left[2\sum_{t=1}^{T} s_{it}s_{jt} + T - \sum_{t=1}^{T} s_{it} - \sum_{t=1}^{T} s_{jt}\right] - (2\hat{\mu}_i\hat{\mu}_j + 1 - \hat{\mu}_i - \hat{\mu}_j) \\
&= \frac{2}{T}\sum_{t=1}^{T} s_{it}s_{jt} - 2\hat{\mu}_i\hat{\mu}_j \\
&= \frac{2}{T}\sum_{t=1}^{T}(s_{it} - \hat{\mu}_i)(s_{jt} - \hat{\mu}_j) \qquad (1.7)
\end{aligned}
$$

$\hat{I}_{mc}$ is proportional to the estimated OLS coefficient $\hat{\beta}$ in the regression of $\bar{s}_{it} = s_{it} - \hat{\mu}_i$ on $\bar{s}_{jt} = s_{jt} - \hat{\mu}_j$.[22] In other words, regressing the first series on the second, –provided that both of them are centred around their mean– is sufficient to implement a test that is consistent with the concordance index approach of the two authors. The test is based on the null $H_0 : \beta = 0$. The problem is that it is highly probable to find serial correlation in the series under study. In such a case, a simple t-ratio test cannot be done and one needs to compute instead t-ratios that are robust to serial correlation.

Correcting for serial correlation

Many methods exist to correct the problem of autocorrelation (and heteroscedasticity), e.g. maximum likelihood estimation, Feasible Generalised Least Squares or GMM. We will not use them because they require some information about the structure of the covariance, which is not available here. Moreover, in this particular test, OLS have to be used for the estimation of β. In that case, the solution is to find an estimator of the appropriate

[22] Indeed, $\hat{\beta} = \sum_{t=1}^{T}(s_{it} - \hat{\mu}_i)(s_{jt} - \hat{\mu}_j) \times \left[\sum_{t=1}^{T}(s_{jt} - \hat{\mu}_j)^2\right]^{-1} = \frac{\hat{I}_{mc}}{2\sigma_K^2}$ where σ_K^2 is a strictly positive constant. Therefore, $\hat{\beta}$ and $\hat{I}_{mc}$ are proportional.

asymptotic covariance matrix. The Newey-West estimators for autocorrelation is the most commonly used tool to compute the robust covariance matrix of $\hat{\beta}$.

Some more recent techniques allow to estimate covariance matrices that are robust to serial correlation and heteroscedasticity at the same time. We will use here one of them, provided by Den Haan[23] for the article of Den Haan & Levin (2000). The idea in this kind of articles is to pre-whiten the errors before computing covariance matrices. As we have just said above, the problem is that the number of lags have to be determined first. In general, one uses a first order VAR. The procedure of Den Haan & Levin estimates a specific lag for each independent variable, using the Akaike's and Schwarz's information criteria.

As a matter of comparison, we also use a correction based on quasi-differences, under the assumption that the serial correlation is simply of order one for all series. In other words, we use the transformation $\bar{s}^*_{it} = \bar{s}_{it} - \hat{\rho}\bar{s}_{it-1}$ where $\hat{\rho}$ is the estimator of the coefficient corresponding to serial correlation. $\hat{\rho}$ is estimated by the Cochrane-Orcutt procedure. If serial correlation actually has this form, the residuals resulting from the regression of $\bar{s}^*_{it}$ onto $\bar{s}^*_{jt}$ are *i.i.d.* which allows to use standard t-ratios. Note that this test rejects the null quite often, thereby suggesting that serial correlation is actually of an order higher than one. However, we are not so much interested in the level of the statistics found than in cross-country comparison. We use this approach as a way to double-check the results found with the DenHaan & Levin correction.

Results

The results of the 306 tests are given in the appendix[24]. A first result of the first test (DenHaan & Levin correction) is that the proportion of rejection is higher in the Euro area than elsewhere. Surprisingly, there are more rejections in the first period than in the second. For example, in the Euro area the null of independence between cycles is rejected 41.8% of the time (46 rejections for 110 tests) whereas in the second period, the proportion falls to 24.5%. But at the same time, the number of rejections for the tests involving non-Euro countries falls dramatically: from 23.6% in the first period to 3.9%. That is, H_0 is rejected 1.77 times more often for the Euro group compared to the other

[23] http://weber.ucsd.edu/~wdenhaan

[24] The procedure of Den Haan allows to choose between the Schwartz criterion and the AIC. As these two were giving exactly the same results, the tests reported have been done with one of them only (AIC).

countries in the first period and 6.23 times more often in the second.

The second test goes in the same direction, with even stronger results since the level of rejection for Euro countries increases slightly (87.3 to 88.2%). At the same time, the level of rejection for tests with others countries decreases by 23.6 points (85.7 to 62.1%).

Business cycles bilateral relations test (Harding & Pagan, 2000a)
Proportion of tests for which the null* has been rejected (percentage)

	Den Haan & Levin correction		Quasi-difference correction	
	First period	Second period	*First period*	Second period
Euro countries	41.82	24.55	87.27	88.18
EU countries	31.87	15.93	87.36	77.47
At least one non-Euro	23.65	3.94	85.71	62.07
At least one non-*EU*	28.13	4.69	84.38	64.06
Average	30.18	11.24	86.20	71.59

**H0*: no statistical relation between two cycles

Looking more in details within the Euro group, we see that a 'core' group appears in the second period. There is more often a link statistically significant in a group composed of Austria, France, Germany, Greece and the Netherlands. The proportion of rejection for the tests involving only countries of this group is 95% in the second period (19 rejections out of 20 tests). Note that no such group was obvious in the first period (only 35% of rejections, that is below the average of the Euro group). For the quasi-difference correction, the rejection for this group increases from 80 to 100% although it is difficult to see any evidence of a core group for this test, given the proportion of rejection.

Everything looks as if the business cycles had become more idiosyncratic across countries, but that this phenomenon was less accentuated within the Euro area. In that case, we could say that the monetary integration has helped creating *some* links between the member countries compared to elsewhere. The potential existence of a core group with strong business cycles links within the Euro group suggests that monetary integration might not be the only channel of business cycles concordance.

It should be noted that the rejection/non-rejection of the hypothesis itself does not give all the information about the dependence relations between two cycles. The non-rejection of the null does not necessarily means its *acceptance*, and it does not inevitably imply that the two series are actually independent.

1.5 Conclusion

This study suggests that the cycles have become more idiosyncratic internationally, by comparing the period before the creation of the EMS (January 1962- February 1979) and the one after (March 1979- December 2004). But at the same time, this phenomenon is less accentuated for the Euro group. Moreover, within it, a smaller group shows some indices of increased business cycles synchronisation. A general conclusion would be that the monetary integration process has been correlated with stronger business cycles links for a *core group* of countries composed of Austria, Belgium, France, Germany, Greece. Luxembourg and the Netherlands. The other countries of the Euro group follow the general movement of the panel towards more idiosyncrasy, but this is less accentuated than for non-Euro countries.

Concerning cycle shapes, no real similarity could be observed between the Euro countries during the first period. Yet in the second period, a greater homogeneity in the shapes of the cycles was present. In particular, cycles lengths are quite similar for Belgium, France, Germany, Italy, Luxembourg and the Netherlands. For the visual aspect of shapes the results are less clearcut, and three groups appear. Among them. French and German cycles are very similar but are different from the others.

The study of time synchronization for the Euro countries has shown that the comovements with the German cycles have decreased (Pearson's coefficient) or increased (Concordance) in the second period. This result is contradictory at first sight. In fact, comovements with the US decrease even more, such that for the two method there is more synchronization towards the German cycle. Only two countries (Finland and Italy) were more synchronised in the second period with the US cycle than with the German one. No clear results were observed for non-Euro countries.

A robust t-ratio test of the dependence between the cycles has also been conducted. It confirms somehow the result of the Pearson's coefficient in the previous part: a) more rejection of the null of independence is observed within the Euro group. but b) the hypothesis is rejected less often in the second period than in the first.

Overall, a core group of Euro countries is observed at every stage of this study, but its composition varies. A constant feature of the second period is that the French and German cycles are very similar. The link with other Euro countries can be strong as well, but depends on the criterion used to measure it. In a sense, France and Germany act as a kernel to which other Euro countries cycles are more or less attracted.

To answer the question raised in the introduction about the optimal currency area, EU countries cycles have followed the general movements through time towards greater independence, which suggests that the EU is not intrinsically an OCA. At the same time, this movement was less important in the Euro group. Besides, as a kernel of more strongly linked countries appears within it, we can suggest that the Euro area might become an OCA in the future.

These results should not hide several methodological limitations. In particular concerning the dating procedure. The fact that different dates from those published by the NBER and the ECRI have been found questions the ability of univariate applications to capture the overall business cycle dates. Talking about the *business* cycles here is perhaps a bit excessive, and it would be better to simply talk about the *industrial sector* cycles. There is a close correspondence between the two in many cases, but this is not enough to generalise the results.

A second problem is directly linked to the previous one and is more general. If the dates vary from one procedure to the other whereas the dating for the US are quite similar, this should tend to confirm the idea exposed by Hamilton (2001) that such algorithms cannot be generalised to any other country. The BB procedure has been designed to reproduce the NBER dating process for the US turning points. The modified versions of this procedure (including the one used here) reach this goal as well. But the apparent sensitivity of the results to the method used for the other countries questions our ability to apply the BB procedure everywhere.

This gives us some perspective for future work. First, it might be of interest to develop a procedure for a *vector* of variables, although we may face a problem in that most macroeconomic series are quarterly. Second, the approach used is only descriptive, and it seems that one cannot go much beyond that with such methods. A proper econometric model would be needed. The next chapters will try to fill this gap.

Chapter 2

The Structure of Trade and Business Cycle Correlations

Abstract

This paper tries to give additional insights on the international transmission of business cycles. We assess the link between intra-industry trade (IIT) and the similarity of business cycles. It is generally agreed that a more similar structure of trade should lead to more coordinated business cycles because they would be more affected by common shocks. We use spectral methods to disentangle information on comovements and information on synchronization. The results show evidence of a positive influence of IIT on both business cycles comovements and synchronization. Moreover, we find a positive influence on cycles that comove and are synchronized at the same time, while there is a negative influence on cycles that comove with a time delay. This last point confirms the idea that IIT induces common supply shocks across countries. In a second part, we distinguish between vertical and horizontal IIT. The results suggest that business cycles correlations are more strongly positively linked to VIIT than HIIT, suggesting that business cycle are correlated provided that the goods exchanged are of the same type but also that they are imperfectly substitutable.

2.1 Introduction

It is often observed that business cycles are coordinated across countries (See e.g. Artis et al., 1997, Artis & Zhang, 1996). We will try here to give additional insights on the reasons of this phenomenon. In particular, we will focus on trade, which seems to be one of the most prevalent channels of transmission of business cycles from one country to another. The idea that an increasingly intensive trade causes greater synchronization of business cycles is not obvious. Indeed, according to the classical theory of international trade, an increase in trade should lead countries to specialize in the production of goods for which they have a comparative advantage. In that case, more intensive trade could entail more idiosyncratic cycles. Krugman (1993) applies this argument to the European Union: economic integration leads to specialization. To the extent that monetary integration prevents stabilization by exchange rate variations, he asserts that the combined effect of economic and monetary integration leads to more severe crisis, using the example of US regions. Besides, a region experiencing high factor mobility will not see its factor prices diminish. Consequently, after a crisis this region will be unable to attract new industries and constantly diverging growth rates across regions will take place. However, for DeGrauwe (1993), labour mobility is unlikely to increase in Europe as much as in the US. Conversely to what Krugman argues, there should be pressures on real wages, allowing countries to catch up with the growth rates of their neighbours. Even though DeGrauwe agrees on the fact that greater specialization should take place, for him the difficulties for Europe should be smaller than what Krugman claims.

Frankel & Rose (1998) empirical findings contrast with Krugman's argument. They observe that larger trade flows are associated with greater business cycles correlation. They suppose that increased trade flows in the EU are the result of monetary and economic integration, and conclude that the EMU is endogenously optimal.

Fontagné & Freudenberg (1999) identify another kind of endogeneity. They find a negative relation between intra-industry trade (IIT) and exchange rate volatility. A consequence would be that monetary integration, by suppressing exchange rate uncertainty, could have raised IIT in the Euro zone. As far as the trade structure is representative of the output structure, cycles should become more synchronized because they would be affected by common shocks. This is the argument of Kenen (1969), who states that more IIT increases the optimality of a monetary zone. Therefore, rather than looking at the influence of trade flows, we will consider the structure of trade between countries.

Reading the two studies of Frankel & Rose and Fontagné & Freudenberg leads to believe in the endogeneity of monetary integration. On one hand, the 'Single Market' and –as argued by FR– monetary integration have boosted trade flows. According to their empirical findings, this has increased the correlation of business cycles. On the other hand, Fontagné & Freudenberg argue that the monetary integration has also led to an increase in the share of IIT in total trade within the euro zone. The addition of these two phenomena should induce greater symmetry of shocks and therefore more business cycles similarity. In other words, the effect of similar structures of trade on business cycles –as a consequence of the latter phenomenon– would be all the more important that there are large trade flows –as a consequence of the former.

As noted by Prasad (1999) and Hoffman (2003), demand and supply shocks have different effects on business cycles. One could imagine two countries that would exchange very little together but that would have a similar structure of production. If supply shocks are dominant to explain international business cycles –as suggested by Ahmed et al., 1993– then it could be that these two countries have coordinated business cycles –see figure 2.2. An example could be an oil price shock that would modify the cost of production of industries in both countries and influence their business cycles in a similar way, even if the two economies are perfectly independent from each other. If, on the contrary, demand shocks dominate, transmission should be proportional to the trade between the two countries. If a country sells a large part of its production to its neighbour and if the latter experiences a recession, the demand for foreign goods will fall –provided the income elasticity of export demand is high– leading to a recession in the first country as well.[1] The joint influence of trade flows and the trade structure might be of importance in the determination of business cycles correlation. The former would foster demand shocks transmission, and the latter supply shocks.

The interest of the question we look at is twofold: first, it can give some insights into the channels of transmission of business cycles and second, it is linked to the debate on the endogeneity of the Optimal Currency Area (OCA) criteria –does monetary integration, and therefore the EMU, lead *de facto* to an optimal monetary zone?

[1]This is argument of the traditional Harrod trade multiplier. The correlation between cross-country outputs will be functions of imports and exports income elasticities.

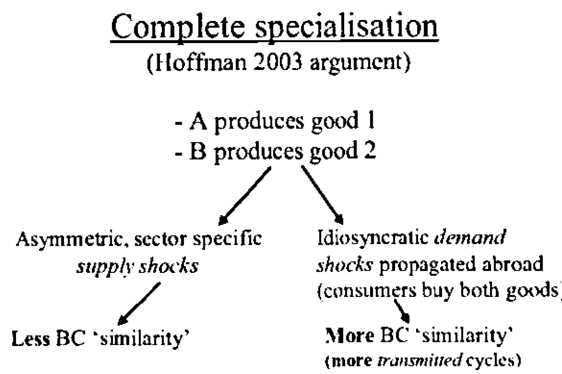

Figure 2.1: One way trade and business cycles

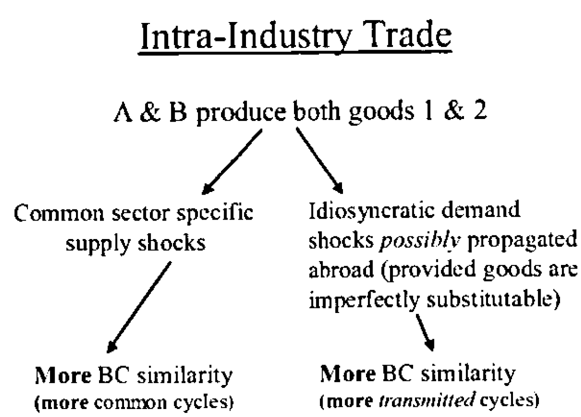

Figure 2.2: IIT and business cycles

This paper tries to contribute to the literature on two respects. In the first part, we use more precise techniques for business cycles similarity than simple contemporaneous correlations. We compute cross-spectral densities for each pair of countries. Such a method allows to disentangle the comovements of two series at a particular frequency and their synchronization. In addition, we use this two dimensional information to assess whether business cycles are subject to common or transmitted shocks.

In the second one, we use a more trivial measure of business cycles similarity, i.e. simple correlations, and we put the emphasis on the measure of the trade structure. Using disaggregated data on trade, we take a more precise measure than the one used in this part of the literature, by distinguishing between vertical and horizontal IIT. We try to identify the type of trade structure that explains these correlations best. As far as I know, no study has yet estimated the influence of vertical and horizontal IIT on business cycles synchronization.

The next part reviews the literature. The two followings are dedicated to empirical studies. The last part concludes.

2.2 A review of the literature

Three main channels of transmission of business cycles are identified in the literature. For instance, Imbs (2001) estimates a simultaneous equation model to find a link between them: specialization, trade flows and financial links. As a result of the increased integration of trade, investment and finance, there should be a higher transmission of business cycles between countries. Indeed, transmission has increased among industrialized countries in the last two decades compared with the two previous ones.

The first channel is the international financial system. In a world of highly integrated finance, the instabilities are rapidly transmitted across countries. RBC models show that in an economy with perfect capital mobility, financial markets allow a sharing of the risks induced by idiosyncratic shocks.[2] The empirical study of Astrubaldi et al. (1996) find that about 40% of shocks to output are smoothed through capital markets in the US. At the same time, Sorensen & Yosha (1996) underline that for European and other OECD countries, risk-sharing rather takes place through national government budget deficit and

[2]among other examples, see Ubide-Querol (1995). A part of this literature also assumes incompleteness of financial markets, e.g. Baxter (1995) and Baxter & Crucini (1995).

corporate savings than through capital markets. Therefore, shocks smoothing are less important among European countries than among American states. Complete risk-sharing implies that individuals do not react to country-specific shocks, but only to common shocks affecting the whole system. Because individuals from every country react to the same shocks, individual business cycles should move –at least partially– in a coordinated way. On the other hand, complete risk-sharing tends to promote specialization –as predicted by Krugman, 1993– with a depressing effect on business cycles synchronization. On this point, Imbs (2004) finds empirically that countries integrated financially tend to be more synchronized even though they are more specialized. Osborn & Sensier (2002) suggest that local interest rates are amongst the main channels of determination of business cycles and that foreign interest rates are an important channel of transmission of these cycles, for all countries except the US.

The second channel is international trade. As noted above, Frankel & Rose (1998, henceforth FR) study the relation between business cycle synchronization and trade. Their idea is that monetary integration processes might endogenously create the conditions for the optimality of a zone. According to standard versions of the OCA theory, different countries would benefit from adopting a common currency if 1) they have coordinated business cycles, 2) there is a high mobility of factors between them and 3) they trade a lot with one another. These three criteria are set outside the system. However, if there was a link between two of the three criteria, and if one rightfully assumed that integration in Europe had reinforced trade linkages, then the optimality of the area would no longer be determined by exogenous factors. It would be endogenous to the system, such that the simple fact of creating a currency area would raise the probability to have an optimal system, at least regarding the three criteria. Of course, this would require that the relation between trade intensity and business cycles comovements is positive.

Several criticisms could be addressed to FR. A first one, pointed out by Imbs (1998), is that there is a difference between trade flows fostered by lowering tariff barriers and trade fostered because of the optimality of a monetary zone. FR claim that because of integration, more trade takes place which leads to more business cycles synchronization. This increases the optimality of the zone. FR conclude that there is endogeneity in such a system. There is the implicit idea that monetary integration will necessarily be linked to deeper economic integration, so that it will increase further trade flows etc. Kenen (2002) also criticizes FR in that it is not because asymmetric shocks are reduced between two economies that their output will necessarily be more correlated. It all depends on the type of shocks prevailing most. This point can be linked to the figures in introduction.

The main criticism, following Kenen (1969), is based on intra-industry trade (IIT). Kenen points out that the production diversification is relevant for the existence of an OCA.[3] The more specialized a country is, the more its business cycles should be independent *vis-à-vis* its neighbors. Diversified economies do not need great variations in their real exchange rates and might therefore be able to afford to share a common currency. FR acknowledge the importance of IIT but they do not integrate it in their estimation, which is done with a measure of total trade intensity –i.e. total bilateral exports or imports– whereas one would need more detailed data showing trade by sectors. This is what Fidrmuc (2004) or Gruben et al. (2002, henceforth GKM) do. Both studies find a positive relation between IIT and business cycles correlations. GKM least squares coefficients are about three times smaller than those of FR by instrumental variables. For the authors, this is due to the inability to find an instrument that would affect trade but not the correlation of cycles. To be valid, instruments must affect the outcome variable, i.e. cycles correlations here, only through the variable instrumented, i.e. trade. In fact, business cycles synchronizations can be affected not only by trade, but also by common monetary policy and by factor mobility. Moreover, those two last elements can be instrumented by the same instruments as trade. This might bias the IV estimation and can explain why the coefficients found by FR are so high. GKM suggest instead to take OLS estimation and to integrate instruments into the equation.

Anderson et al. (1999) use spectral methods and the multivariate version of the Beveridge-Nelson decomposition of Stock & Watson (BNSW) to evaluate the link between international trade and business cycle synchronizations. Using a test for the existence of common cycles, they find that countries that were relatively open during the past three decades were more likely to have common business cycles with their major economic partners. An interesting fact is that they find a much weaker relation between openness and the coherence of HP-filtered series. In general, their estimations made with the coherence measure are less intuitive than those obtained with the BNSW decomposition. For them,

[3] Kenen (2002) emphasizes on the fact that the degree of diversification must *not* be the only criterion of optimality. He claims that Kenen (1969) might have put too much emphasis on this point.

this result is due to the fact that coherence is largely influenced by contemporaneous correlations[4].

The problem at this point is that there is still a missing part between monetary integration and the degree of specialization. For the 'endogeneity argument' to be valid, there should be a negative relation between these two elements. The contribution of Fontagné & Freudenberg (1999) suggests that the European monetary integration process induces higher IIT. For two reasons. 1/ A stabilization of exchange rates reduces uncertainty. Risk-averse firms do not fear a drop in profits due to an unexpected exchange rate appreciation, which would entail a decline in foreign demand. Consequently, they have less incentives to produce in the country relatively specialized in their industry. Leaving aside considerations about transaction costs, the result is more homogeneous production structures across countries. *In fine*, this leads to higher IIT and therefore more symmetric shocks. Said differently, a positive demand shock for one particular good will have the following consequences under ***flexible*** exchange rates: the country relatively specialized in this good will see its currency appreciate and the expected profits of the firms that produce this good will increase. This will drive foreign firms to relocate in this country, increasing specialization further. Note that this argument of Fontagné & Freudenberg contradicts that of Krugman (1993). The latter focuses on reduced transaction costs which would incite firms to 'agglomerate' and locate in places where other firms of the same type have already settled. This leads to greater specialization. 2/ The impact of exchange rates on trade may differ with the nature of trade. That is, trade flows between completely specialized countries may not be too highly affected by variations in exchange rates. On the contrary, if there is already a high degree of intra-industry trade, the elasticity of demand should be high and trade should vary greatly with exchange rates. Consequently, monetary integration should have a larger impact on trade flows when a high degree of IIT takes place –as it is the case in the EU– than when production is specialized.

To sum up, the idea of FR is that an area might become optimal *ex-post* even if it fails to be one ex-ante. Conversely, economists such as Krugman (1993) claim that an area might fail to be optimal ex-post even if it passes the criteria *ex-ante*.

There are also skeptical views about trade as a channel for business cycles transmission.

[4]This may be due to the fact that stochastic trends have non-null spectral density at any frequency. See section 2.3.1 on this point.

The main conclusion of the study of Flandreau & Maurel (2002) on historical data is that trade may not be so important in the transmission of business cycles. Besides, they find that monetary integration is associated with more intensive capital exchanges, and more portfolio diversification. Their main conclusion is that increased comovements of output fluctuations are associated with more intensive and *specialized* trade. This conclusion is rather unusual because it implies that even if monetary integration leads to a greater specialization of output, business cycles can be more synchronized. Consequently, trade alone might not be so important in the process of increasing comovements.

Imbs (1998) criticizes the approach of Frankel & Rose and in particular their choice to use geography or common language as instruments. He claims that the essential question is to assess the cost of giving-up an independent monetary policy. For this purpose, one must measure business cycles synchronizations, abstracting from the effect of monetary policy. Indeed, there is a difference between a rise in trade due to an OCA and a rise in trade due to lower tariff barriers. He uses correlations of bilateral *total factor productivities* and bilateral labour productivities. Then he estimates the relation between these measures and trade, and concludes that trade does not have a significant influence on the comovements of business cycles. Note however that this is not the case for European countries. Imbs (1999) uses a monopolistic competition model with two countries and shows that there are two asymmetric equilibria –where most of the heterogeneous good is produced in one country– and one symmetric equilibrium. The first two are unstable whereas the latter is stable, so that the global economy tends to be symmetric asymptotically, which leads to symmetric cycles as well.

The multi-sector model of Ambler et al. (2002) manages to generate positive international transmission of the business cycles, which is an improvement upon other RBC models. They find that the mechanism leading to positive correlation of the cycles is not trade in intermediate goods. It is rather the existence of several sectors: technological shocks in one sector of a particular country attract labor and capital from abroad and also from the other sectors of the country. This tends to create positive comovements. This finding is all the more interesting that there is simultaneous exports and imports of the same type of goods in all sectors. In other words, they manage to reproduce some intra-industry trade. Therefore, the model of Ambler et al. indirectly approves the idea that cycles are due to IIT –i.e. here the existence of many sectors coupled with simultaneous exchanges in each sector– rather than trade flows.

To sum up, it seems that the literature agrees on the fact that it is more the degree of (non-)specialization of a country that determines business cycles comovements than its degree of openness.

2.3 Distinguishing common and transmitted cycles

2.3.1 Methodology

Two types of estimations will be conducted in this first empirical part. We look at the relation between IIT and :

a) comovements strictly speaking –i.e. to what extent are two business cycles dominated by the same frequencies?

b) synchronizations –i.e. to what extent will two elements with the same frequency lead or lag each other?

The data used is the monthly, seasonally adjusted index of industrial production, provided by the OECD. The data set comprises 16 countries and the sample starts in January 1967 and ends in January 2001[5]. We use industrial production as a proxy for output, following Artis et al. (1997) or Massman & Mitchell (2002). Data on trade comes from the CHELEM database of the CEPII. It gives bilateral trade flows for 71 products. Its frequency is annual. Measures of bilateral distances in kilometers between capital cities are also provided by the CEPII.

Measuring business cycles 'similarity'

Defining business cycles Recall that in the empirical business cycles literature, different definitions coexist under the word 'cycle'. We consider two of the most widely used. A cycle can be defined as the stationary remaining part once the trend –i.e. the non-stationary part– has been removed, e.g. Engle & Kozicki (1993), Vahid & Engle (1993) (definition 1). It can also be the set of elements having a frequency lying within a

[5]The countries are: Austria, Denmark, Finland, France, Germany, Greece, Italy, Japan, the Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, United-Kingdom and the United States. Originally, Belgium and Luxembourg were present in the dataset. They have been suppressed because of a lack of data for trade.

predefined band[6], e.g. Baxter & King (1999) (definition 2).

The series potentially include stochastic trends which are non-stationary elements by definition. The problem is that these elements have non-null spectral densities at any frequency[7]. For this reason, the cycles will be computed twice, over series in level and over filtered, stationary series. Measuring the cycles according to definition 2 *stricto sensu* might leave some non-stationary elements in the cycle.[8] This is the case when cycles are extracted by means of a linear filter if the weights do not sum up to zero[9]. In that case, we have a conflict between the two definitions:

1/ the cycle produced by selecting elements from a band of frequencies contains parts of the trend, i.e. we face a leakage problem analogous to that pointed out by Harvey & Jaeger for the HP filter, and

2/ at the same time, removing the non-stationary part implies a modification of the elements lying within the business cycles frequency band, which is precisely what the second definition wants to avoid.

Cycles are first computed from the raw series in level and second from pre-filtered series –first differencing[10] and Baxter-King's bandpass filter have been used for pre-filtering. The first approach respects the 'band of frequencies' definition in that it does not modify the elements of the band but might be in conflict with the definition of a cycle as an element 'without trend'. The second approach has the advantage of not being in contradiction with the second definition. Said differently, if a pair of series does not share a common stochastic trend, one should look at the first measure, but they do share it, there will be a conflict between the two measures and one might prefer the second one. The interest of computing two types of cycles is to see how differences between them induce diverging estimations. We will see that estimations diverge indeed and sometimes quite a lot. We will tend to follow the literature and to favour the approach based on prefiltering data prior to

[6]Usually in business cycles literature, one eliminates all elements with a frequency different from ω^*, $\frac{2\pi}{p1} \leq \omega^* \leq \frac{2\pi}{p2}$, where $p1 = 32$, $p2 = 6$ for quarterly data.

[7]A proof of this well-know result is available upon request.

[8]This kind of argument has been used concerning 'spurious cycles' induced by the HP filter. See Harvey & Jaeger (1993).

[9]An example of this property can be found in Baxter & King (1999). BK filter weights are enforced to sum to zero in order to have trend removing properties.

[10]It is worth pointing out, and well-known in the literature, that first differencing produces series biased towards high frequencies. We use here this measure as a matter of comparison.

analysis. The potential existence of common stochastic trends in the first type of cycles might induce oversized correlations which might bias the estimations. The inconvenient is that this approach does not fit with that purely based on frequencies.

Two techniques are used for the comovements of business cycles. Both are based on the cross-spectral density $F_{xy}(\omega)$. In continuous time and doubly infinite sample, it is the Fourier transform of the cross-covariance function $C_{xy}(\tau) = E[x_t y_{t+\tau}]$ where $E[.]$ denotes expectations. It is defined as

$$F_{xy}(\omega) = \frac{1}{2\pi}\int_{-\infty}^{\infty} e^{-i\omega\tau} C_{xy}(\tau) d\tau \tag{2.1}$$

where x and y denote the time series and ω the frequency. The first one is the spectral coherence. See also Anderson et al. (1999) for a similar use of this tool. Coherences return the intensity of the comovements between two series at a given frequency. The second one is the phase lag and it measures the time delay. If two cycles comove perfectly, two situations might happen: in the first one, the series are perfectly synchronized, and we could say that they are fully determined by common shocks. In the second situation, one series leads the other in time, and we might suspect that shocks are transmitted from one series to the other.

Measuring comovements: average coherence We first take the spectral coherence $\rho_{xy}(\omega)$ between the series for each frequency[11], assuming that the frequencies are independent between them and *a fortiori* from one country to another.

$$\rho_{xy}(\omega) = \frac{F_{xy}(\omega)}{(F_x(\omega)F_y(\omega))^{1/2}} \tag{2.2}$$

The elements lying at undesired frequencies –i.e. out of the business cycles range– are suppressed and the remaining elements are averaged. This entails that two series will have a high average coherence if they are dominated by the same frequencies. One could interpret the coherence as the correlation between the series x and y *for one particular frequency.* Eq.(2.1) shows that a time series can be decomposed into an infinite sum of elements over the interval $[-\pi, \pi]$. Therefore, one would need to add up the coherences over different frequencies to be able to interpret the information in the time domain. However, such a sum *could no longer be interpreted as a correlation.* Correlations have the form $\frac{\int A}{(\int B \int C)^{1/2}}$

[11]More exactly, we use the modulus of the coherence, since it is a complex number: $|\rho_{xy}(\omega)| = \sqrt{Im(\rho_{xy}(\omega))^2 + Re(\rho_{xy}(\omega))^2}$

whereas[12] the sum of coherence has the form: $\int \frac{A}{(BC)^{1/2}}$. This might explain the fact that Anderson et al. (2002) find a strong correspondence, but not a perfect equality, between the average of coherence of two series and their correlation. Taking the average value of the coherence (assuming to simplify that we are only interested in the frequencies lower than $\overline{\omega}$) we get,

$$\frac{1}{2\overline{\omega}}\int_{-\overline{\omega}}^{\overline{\omega}} \rho_{xy}(d\omega) = \frac{1}{2\overline{\omega}}\int_{-\overline{\omega}}^{\overline{\omega}} \frac{F_{xy}(d\omega)}{(F_x(d\omega)F_y(d\omega))^{1/2}} \tag{2.3}$$

One advantage of using coherence instead of simple correlations is that it allows us to take the information at some particular frequencies only. Correlations take instead the information from the whole spectrum, and pre-filtering is required to get the cycles correlations. Thus, taking the average coherence avoids the possible distortions of the cycles due to filtering[13]. Canova (1998) pointed out that cycles change with the detrending method used. See also Guay & St-Amant (1996) for a particular discussion of HP and Baxter-King filters.

It can happen that two series have a high coherence even if one leads the other. Indeed, there can be a peak in their cross-spectrum at a particular frequency even though the elements of the series at this frequency are not perfectly synchronized. To understand this point, take a simple example. Let $x_t = \sin(\lambda t)$ and $y_t = \sin(\lambda t + d)$ where λ and d are constants. The whole bulk of the cross-spectrum between x_t and y_t is concentrated at frequency λ even if y_t leads x_t by d periods. This is illustrated in the plot below.

On the other hand, the larger d is, the lower the contemporaneous correlation will be. In other words, contemporaneous correlation is unable to take lagged comovements into account. We can say that coherence measures the correlation between two series in the

[12]Covariances and variances can be expressed as $cov(x,y) = C_{xy}(0) = \int_{-\pi}^{\pi} F_{xy}(d\omega)$ and $var(m) = C_m(0) = \int_{-\pi}^{\pi} F_m(d\omega)$ for $m = x, y$ such that correlations are
$\zeta_{xy} = \frac{cov(x,y)}{\sigma_x\sigma_y} = \frac{\int_{-\pi}^{\pi} F_{xy}(d\omega)}{\left(\int_{-\pi}^{\pi} F_x(d\omega)\int_{-\pi}^{\pi} F_y(d\omega)\right)^{1/2}}$

[13]However, in order to estimate the cross-spectrum, one needs to use a so-called *window*, which smoothes somehow the periodogram –that is, the 'raw' estimate of the spectral density. Consequently, it takes for each frequency some information from the adjacent frequencies. At the opposite, taking the periodogram without using a window would induce too much variability. It is therefore necessary to select the window appropriately so that the distortions that it involves do not exceed those of the filter. In other words, one has to ensure that the 'side lobes' problem of the filter is greater than the 'window leakage' of the cross-spectrum. I have used here the Parzen window, with a lag parameter of approximately $T/2$. This is higher than in the usual estimation of spectrums where the lag parameter is often $T/20$. This should reduce the window leakage. Recall that a lag parameter equal to the sample length T gives the periodogram.

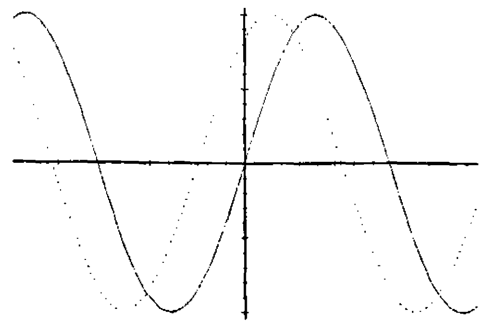

Figure 2.3: $x_t = \sin(\lambda t)$ (solid) and $y_t = \sin(\lambda t + d)$ (dotted)

frequency domain only, such that the time synchronization does not influence it.

Measuring synchronizations: phase lags The function that allows to measure the time concordance is the *phase lag* function $\varphi_{xy}(\omega)$.

$$\varphi_{xy}(\omega) = \arctan\left[\frac{\mathrm{Im}(F_{xy}(\omega))}{\mathrm{Re}(F_{xy}(\omega))}\right] \tag{2.4}$$

We will take the absolute value of this term. $\varphi_{xy}(\omega)$ can be regarded as a measure of the time delay between two series. A phase lag equal to zero at any frequency implies perfectly synchronized series. It is the argument of the cross-spectrum $F_{xy}(\omega)$ (recall that in general $F_{xy}(\omega)$ is complex). Intuitively, the argument of a complex number z ($\arg(z)$) measures the angle between the vector $(0, z)$ and the real axis in the complex plane. Note that the larger $\mathrm{Re}(z)$ relative to $\mathrm{Im}(z)$, the smaller $\arg(z)$.[14] Note also that in order to average the function, the same kind of procedure as for coherence is used.

This ability of spectral analysis to distinguish between these two aspects of comovements allows to think in terms of transmission of the cycles, unlike contemporaneous correlation. Indeed, two series might have a high coherence and a non-null phase lag at some frequency ω. This means that the elements of the two series corresponding to ω will

[14] Two elements might give an intuition of why the phase lag function is a measure of time delay between two series. First, it can be shown that a real cross-spectrum implies that the cross-covariance function is even, so that a null phase-shift entails perfect synchronizations. Second, a cross-spectrum composed of an imaginary part only implies that the two series are completely uncorrelated. Proofs of these statements have not been reproduced here but are available upon request.

comove but not in a synchronized way. This case can be regarded as a proxy for a measure of transmission of business cycles. On the other hand, a high coherence combined with a null phase lag might be an indication that the two series are affected by common shocks.

Business cycles transmission vs. common business cycles

We also use two measures derived from the previous ones in order to capture *transmission* and *common* business cycles relations. We will consider that there is a transmission mechanism when two cycles have a high coherence and a significant phase lag between them. In order to capture bilateral business cycles relations that exhibit such a feature, the following measure is used in the estimations of the next part.

$$Transmission_{xy}(\omega) = \rho_{xy}(\omega).|\varphi_{xy}(\omega)| \tag{2.5}$$

Where $\rho_{xy}(\omega)$ denotes the modulus of coherence. Similarly, we will consider that two cycles have a common cycle if they comove and are synchronized, i.e. if there is a high coherence and a small phase lag between them. A proxy for this relation could be

$$CommonCycle_{xy}(\omega) = \rho_{xy}(\omega)/|\varphi_{xy}(\omega)| \tag{2.6}$$

Measures of trade

Two measures of trade specialization are considered in the following.

- In order to measure the similarity in the structure of trade, we use the Grubel & Lloyd Index (GLI), defined as:

$$\begin{aligned} GLI_{i,j,t} &= \frac{\sum_k (X_{ijtk} + M_{ijtk}) - \sum_k |X_{ijtk} - M_{ijtk}|}{\sum_k (X_{ijtk} + M_{ijtk})} && (2.7) \\ &= 1 - \frac{\sum_k |X_{ijtk} - M_{ijtk}|}{\sum_k (X_{ijtk} + M_{ijtk})} && (2.8) \end{aligned}$$

  Where X_{ijtk} and M_{ijtk} denote the exports and the imports, respectively, from country i to country j at time t for good k.

  Note that we follow GKM in that "reported data is more reliable for imports than exports", so that we use M_{ij} instead of X_{ji}.

- Some countries may have a high IIT even though the share of their exchanges is very small compared to the total amount they trade with the rest of the world. In this case, it is less likely that the business cycles of these two countries would interact. This is why we need to weight the GL index so that countries that exchange a lot together have a greater influence in the estimation. We take FR's trade intensity as the weight. It is given by :

$$TI_{ijt} = \frac{X_{ijt} + M_{ijt}}{X_{i.t} + M_{i.t} + X_{.jt} + M_{.jt}} \tag{2.9}$$

where $(X_{i.t} + M_{i.t})$ and $(X_{.jt} + M_{.jt})$ are the total exchanges of countries and i and j, respectively. FR use another measure of trade intensity, where bilateral trade is normalized by GDP instead of total trade. As this measure seemed to suit the data less, we did not consider it here.

Estimations

Our approach is similar to the one of FR. Their paper is an important contribution in the empirical literature on business cycles transmission. However, we shall not forget the criticisms levelled at their methodology while interpreting our results. This is why we also implement the estimation procedure of Gruben, Koo & Millis (2002).

We use a panel of 16 countries, which makes 120 pairs of countries. Four periods of equal length are considered, corresponding roughly to a decade each[15]. This makes a total of 480 observations, but in practice, the total number of observations is 423[16]. For each pair and each period (of about 120 months each), the average coherence and phase lags were computed. Because we suspect that economic integration might have an influence on the similarity of business cycles as well as the structure of trade, we have also made a distinction between a group of EU countries and a group composed of pairs including at least one non-EU country.

Following FR, we use a Two Stage Least Square (TSLS) estimation and we take the same instruments as the authors in the first step of the procedure. Note that estimations were computed on pooled data.

[15]The four periods are: Jan:67-Jun:75, Jul:75-Dec:83, Jan:84-Jun:92, Jul:92-Jan:2001

[16]Because of missing values (e.g. series for industrial production of Spain in the first period or Denmark in the first two periods).

The main equation[17] is:

$$Co_{ijt} = \alpha + \beta IIT_{ijt} + u_{ijt} \tag{2.10}$$

Where Co_{ijt} denotes the measure of business cycles similarity –average coherence, absolute phase lag[18] or the measures for transmission and common cycles explained in paragraph 2.3.1. IIT_{ijt} is either GLI_{ijt} or $TI_{ijt}.GLI_{ijt}$and u_{ijt} is the error term. We are interested here in the coefficient β, which shows whether there is a relationship between IIT and business cycles comovements.

The first step of the TSLS estimation is to 'instrument' the independent variable of the main equation. The instruments used are the logarithms of distances between countries and two dummies equal to one when two countries are adjacent and when the same language is spoken[19]. These variables are the same as those of FR to explain trade intensity. They are often used because of their ability to explain international trade.[20] We have taken them because intuitively countries that are close geographically are likely to have similar structures of production and therefore a high IIT. The first step equation is:

$$IIT_{ijt} = \gamma + \delta \ln(dist_{ij}) + \epsilon border_{ij} + \zeta lang_{ij} + v_{ijt} \tag{2.11}$$

where $dist_{ij}$ is the distance between i and j, $border_{ij}$ and $lang_{ij}$ are the dummies for adjacency and common language, respectively. v_{ijt} is the error term and δ, ϵ, ζ are the coefficients. Instrumenting trade comes down to keep the part of this variable explained by

[17]This is the baseline equation. We will also include time and EU group dummy variables in it. These are not represented for presentation purpose.

[18]The intrinsic idea is that a more similar trade structure should lower the absolute value of the average phase-lag function, i.e. it should increase the average synchronisation.

[19]Note that for the dummy variables, two exceptions have been done: the pair Denmark/Sweden was determined as adjacent, even if they are separated by the sea. In effect, they are very close geographically and are linked by a bridge. The pair Finland/Sweden was considered to have a common language because Swedish is the second official language of Finland.

[20]FR relate these instruments to gravity equations. However, this first step equation is not a gravity equation. The reason it that it excludes any 'mass' variable, such as output. Indeed, the exogenous variables are constant over time, such that the predicted value would be a constant as well. See Fidrmuc & Fidrmuc (2003) for a use of proper gravity models. Note that a simple form of the gravity equation is: $Trade_{ij} = \alpha.output_i^{\beta}.output_j^{\gamma}/distance_{ij}^{\delta}$. The problem is that output cannot be regarded as an instrument for the obvious reason that the LHS of the main equation measures comovements in industrial production series, which implies that the instruments and the LHS would be correlated. In addition, gravity equations are designed to explain trade flows, not IIT.

the instruments and to suppress the remaining part (which is implicitly the part explained by common monetary policies). This way, the second stage allows to see the link between the dependent variable and trade without the biasing influence of common policies.

We have seen above that for GKM, IV estimated coefficients might be oversized. Following their argument we estimate also the following equations, which correspond to eq.(6) of the authors.

$$Co_{ijt} = \alpha + \beta IIT_{ijt} + \delta \ln(dist_{ij}) + \epsilon border_{ij} + \zeta lang_{ij} + u_{ijt} \tag{2.12}$$

$$\begin{aligned} Co_{ijt} &= \alpha + \beta_1 TI_{ijt}.IIT_{ijt} + \beta_2 TI_{ijt}.(1 - IIT_{ijt}) & (2.13) \\ & +\delta \ln(dist_{ij}) + \epsilon border_{ij} + \zeta lang_{ij} + u_{ijt} & (2.14) \end{aligned}$$

Note the presence in eq.(2.13) of a variable corresponding to trade specialization. $TI_{ijt}.(1-IIT_{ijt})$. (2.13) is in fact a decomposition of the equation of FR above. $IIT.TI$ can also be interpreted as a weighted measure of intra-trade.

It is suspected, as in FR, that being in the fixed exchange rate system reinforces the transmission of the cycles[21], as it weakens the independence of monetary policies, which become more similar across countries. To take this argument into account, regressions have also been conducted with a dummy variable capturing exchange rates agreements between countries such as the one that occured trough the *European Monetary System* (EMS). The results were almost not affected, such that we did not reproduce them here.

2.3.2 Results

The results for the first stage equations are displayed in table 3.1. As in FR, the estimations reveal an adjacency effect as well as an effect of distance. Results for EU and non-EU countries have been displayed in appendix A.2.1. A surprising result is that common language is not significant for EU countries, while it is for non-EU countries. The first column displays the coefficient of the equation for trade intensity. The associated R^2 is about the same as those for IIT. This means that the instruments predict as well IIT than trade intensity. Consequently, these instruments can be seen as valuable.

[21] An illustration of this can be found in Massman & Mitchell (1995). They note that the correlation of the UK with the rest of Europe was strongest when exchange rates were fixed – during the short period when the UK entered into the European Monetary System, in 1990-92.

Estimated coefficients from the first stage

	Trade Intensity	*IIT*	*weighted IIT*
Com. Language	0.40	10.78 ***	0.38 **
Distance	-0.32 ***	-6.97 ***	-0.25 ***
Com. Borders	1.80 ***	9.29 ***	1.09 ***
R-squared	0.32	0.32	0.39

Rows: independent variables
Columns: dependent variables
Coefficients are multiplied by 100
*** : 1% significance level
** : 5%

Table 2.1: IIT regressed on instruments

Comovements and phase lags

Tables 2.2 and 2.3 display business cycles comovements using the methodology of FR. The first two lines report for information the measures of correlations on filtered series ('*Corrdiff*' and '*Corrbk*') as used by FR and GKM. The coefficients are all positive and most are strongly significant. In order to assess the robustness of the results, different specifications of the equations have been tried by adding fixed time effects and a dummy equal to one when the two countries belong to the European Union. A noticeable feature is that the coefficients are fairly stable across specifications. Second, time effects are almost always significant, indicating an evolution in time. However, time effects do not alter the significance nor the level of coefficients. A last point to be noticed is that adding the EU dummy affects the coefficients significance. Two of the coefficients associated with coherence are not significant anymore. Besides, the corresponding Wald tests reject the null, indicating a significant effect of this dummy. It could be seen as an evidence that the simple fact of being in the EU leads to higher business cycles comovements. This weakens the relation between comovements and IIT. In order to outline this result, separated estimations between EU and non-EU countries are given in appendix A.2.1, confirming the fact that different relations between EU and non-EU countries are found.

The results for table 2.3 give approximately the same indications, with a positive relation between comovements and IIT and with a strong 'EU effect', which leads some coefficients to be non significant.

The estimations for the OLS approach of GKM are displayed in appendix A.2.1. They give approximately the same results. The coefficients differ from the ones found with IV estimations but are in the same range.

IV Regressions of Comovements on IIT

	models associated to dummies: none	time	EU group	time & EU
Dept. variables				
- IIT Coefficients				
Corrdiff	0.18 ***	0.19 ***	0.16 **	0.16 **
Corrbk	0.47 ***	0.50 ***	0.48 ***	0.48 ***
Coher	0.24 **	0.28 **	0.18	0.20
Coherdiff	0.16 **	0.20 **	0.16	0.17
Coherbk	0.29 ***	0.32 ***	0.28 **	0.32 ***
*- Wald tests**				
Corrdiff	NA	***		**
Corrbk	NA	**		***
Coher	NA	***	***	
Coherdiff	NA	***	***	
Coherbk	NA	***	***	***

*Level of rejection for Wald tests with H0: no joint significance of the dummies included in the equation
Corr diff: correlation on differenced series
Corr BK: correlation on BK-filtered series
Coher: coherence averaged over a range of frequencies
Coher diff: same as 'coherence' on differenced series.
Coher BK: same as 'coherence' on BK-filtered series.

Table 2.2: Comovements and IIT

Tables 2.4 examines the relation between ***absolute phase lags*** and IIT. If countries have more similar structures of trade, they should respond more similarly to the same shocks, which should induce more synchronized business cycles. In short, there should be a negative relation between the two variables. This is the case for the second type of business cycles, that is the ones computed on stationarized series. In that case, more IIT is associated with more synchronized cycles –or less 'time-delayed' cycles. There is an important difference with the first type of cycles, for which the coefficients are positive. This is counter-intuitive and seems to confirm our prior doubts concerning this measure.

As for table 2.2, there is a strong time effect as well as an effect of being in the EU. In addition, one of the coefficients associated to the specification with the EU dummy is not significant, confirming the previous finding that distinguishing EU from non-EU countries might weaken the relation between IIT and business cycles.

The GKM estimations in appendix A.2.1 reveal the same kind of findings. There is an important difference between the two types of cycles. In addition, the positive coefficients

IV Regressions of Comovements on weighted IIT				
	models associated to dummies:			
	none	time	EU group	time & EU
Dept. variables				
- IIT Coefficients				
Corrdiff	3.17 ***	3.27 **	2.78	2.77
Corrbk	9.49 ***	9.78 ***	9.59 ***	9.71 ***
Coher	5.63 ***	6.18 ***	5.19 **	5.83 ***
Coherdiff	3.52 ***	4.10 ***	3.58 **	3.84 **
Coherbk	5.12 ***	5.45 ***	4.81	5.37 ***
- Wald tests				
Corrdiff	NA		***	
Corrbk	NA	***	***	***
Coher	NA	***	***	***
Coherdiff	NA	***	***	**
Coherbk	NA	***	***	***

Table 2.3: Comovements and weighted IIT

associated with the first method are not always significant while the negative ones associated with the second method are strongly significant. As before, the Wald tests reject the null hypothesis almost everywhere, showing a time and an EU effect.

To summarize, the results underline that differences between the two ways of defining business cycles have an important impact on estimations. The first one gives counter-intuitive results in the relation of (non-)synchronicity with IIT. The second one is more in line with theory. This method seems to confirm the findings of FR and GKM. There is a positive and significant relation between the structure of trade and comovements, and a negative one with time delays. However, the results are not as clear for the specification that includes a dummy equal to one when two countries belong to the EU. It seems that belonging to the EU is sufficient to increase the level of comovements and synchronization and to weaken the influence of IIT. Consequently, one might suspect the influence of other variables than the trade structure in the determination of business cycles similarity, in particular variables linked to economic integration.

IV Regressions of phase lags on IIT

	models associated to dummies: none	time	EU group	time & EU
Dept. variables				
- IIT Coefficients				
Aphlag	0.79 **	0.74 *	1.25 **	1.26 ***
Aphlagdiff	-0.84 **	-0.93 ***	-0.65	-0.69 *
Aphlagbk	-1.67 ***	-1.74 ***	-1.75 ***	-1.73 ***
- Wald tests				
Aphlag	NA	***	***	***
Aphagdiff	NA	***	***	*
Aphlagbk	NA	***	***	***

Aphlag: absolute value of phase lags.
Aphlag diff: absolute value of phase lags computed on differenced series.
Aphlag: absolute value of phase lags computed on BK-filtered series.

Table 2.4: Phase lags and IIT

IV Regressions of phase lags on weighted IIT

	models associated to dummies: none	time	EU group	time & EU
Dept. variables				
- IIT Coefficients				
Aphlag	10.61 *	9.59	14.47 *	13.77 *
Aphlagdiff	-16.44 ***	-17.58 ***	-14.13 **	-14.97 **
Aphlagbk	-29.28 ***	-29.83 ***	-29.13 ***	-28.98 ***
- Wald tests				
Aphlag	NA	***	***	*
Aphagdiff	NA	***	***	**
Aphlagbk	NA	***	***	***

Table 2.5: Phase lags and weighted IIT

Measures of 'transmission' and 'common cycles'

IV Regressions of the Common cycles measure on IIT

	models associated to dummies: none	time	EU group	time & EU
Dept. variables				
- IIT Coefficients				
Coms	-0.89	-0.64	-1.91	-1.63
Comsdiff	2.89 ***	3.21 ***	2.75 **	2.82 **
Comsbk	3.41 ***	3.61 ***	3.56 ***	3.64 ***
*- Wald tests**				
Coms	NA	***	***	
Comsdiff	NA	***	***	**
Comsbk	NA	***	***	***

Table 2.6: Common cycles and IIT

IV Regressions of the Common cycles measure on weighted IIT

	models associated to dummies: none	time	EU group	time & EU
Dept. variables				
- IIT Coefficients				
Coms	-5.40	-1.09	-13.12	-6.69
Comsdiff	56.45 ***	60.44 ***	53.88 ***	55.90 ***
Comsbk	60.69 ***	62.80 ***	60.20 ***	61.68 ***
- Wald tests				
Coms	NA	***	***	
Comsdiff	NA	***	**	***
Comsbk	NA	***	***	***

Table 2.7: Common cycles and weighted IIT

Tables 2.6 and 2.7 show the IV estimations of the *common cycles* measure onto IIT.[22] Once again the values for the first line –corresponding to the business cycles measured on raw series– are different from the two others –corresponding to business cycles evaluated on pre-filtered series. There is a non-significant relation between IIT and the 'common cycle measure' for the first method and a positive and significant one for the second method.

[22]The variables for common cycles and transmission have been log transformed in order to avoid the influence of extreme values, due to the product involved in the computation.

As stated above, we shall rather trust the second method because of the potential biasing effect of common stochastic trends. The results from this second measure show that a more similar structure of trade induces more comovements that are associated with synchronization. This is in line with the idea that two countries that exchange the same goods could be seen as having the same structure of production and consequently as being affected by similar supply shocks.

The results for the GKM approach show more similar results between the two business cycles measures. The coefficients are still positive, although lower than for IV estimations. Once again, the inclusion of an EU effect makes that some coefficients are not significantly different from zero. See in particular table A.6 in appendix A.2.1 on weighted IIT.

Tables 2.8 and 2.9 display the estimations with the 'transmission' measure as the dependent variable. The results differ largely between the two definitions of business cycles. The first method returns positive coefficients while the second returns negative ones. The GKM OLS approach produces coefficients that are rarely significant and even never significant for regressions on weighted IIT (table 2.9).

IV Regressions of the Transmission measure on IIT

	models associated to dummies: none	time	EU group	time & EU
Dept. variables				
- IIT Coefficients				
Trans	1.76 ***	1.70 ***	2.47 ***	2.31 ***
Transdiff	-1.86 ***	-1.88 ***	-1.76 **	-1.80 **
Transbk	-1.67 ***	-1.67 ***	-1.65 **	-1.55 **
- Wald tests				
Trans	NA	***	***	***
Transdiff	NA	**	***	**
Transbk	NA	***	***	**

Table 2.8: 'Transmitted' cycles and IIT

It is worth comparing tables 2.6 to 2.9. Focusing on the second type of business cycles ('...diff' and '...bk'), we have seen that IIT is positively correlated with the comovements that are associated with more synchronization (i.e. 'common cycles'). At the same time, IIT is negatively correlated with the comovements that are associated with time delays (i.e. 'transmitted cycles'). We see that even when the coefficients of table 2.8 (or 2.9)

IV Regressions of the Transmission measure on weighted IIT				
	models associated to dummies:			
	none	time	EU group	time & EU
Dept. variables				
- IIT Coefficients				
Trans	25.98 **	24.46 **	31.34 **	28.05 **
Transdiff	-33.75 **	-33.58 **	-31.49 **	-31.87 **
Transbk	-28.38 **	-27.69 **	-26.69 **	-25.22 **
- Wald tests				
Trans	NA	***	***	**
Transdiff	NA	**	***	**
Transbk	NA	***	***	**

Table 2.9: 'Transmitted' cycles and weighted IIT

are significant, they remain lower in absolute value than the ones of table 2.6 (or 2.7). In other words, the effect of IIT on 'common cycles' should be more important than the effect on 'transmitted cycles'. Globally, a more similar structure of trade should induce more similar business cycles through common (supply) shocks.[23]

What about the endogeneity of the OCA criteria?

The results above can be useful for answering the questions of why and how business cycles are transmitted internationally. However, the interpretation in the perspective of the OCA theory is less easy. FR suggest that the positive relation between trade and business cycles correlations can be seen as evidence that the EU –and following this argument, any free trade area– creates by itself the conditions for being an OCA. Indeed, to the extent that the European integration process has accentuated trade flows, this should have deepened the links between European cycles, and therefore should have reinforced the probability for the EU to be an optimal monetary zone.

Can we use the same argument with intra-trade? Fontagné & Freudenberg (1999) claim that the monetary integration creates endogenously more IIT. However, two points have to be noticed. They weaken the idea of endogeneity of the OCA criteria for the EU.

[23]rather than demand shocks. See figure 2.2 above. However, the next part suggests that demand shocks also play a role in business cycles correlations: non-substitutable similar goods have more impact on correlations than substitutable ones .

First, a positive and significant relation is found between IIT and trade intensity for our data[24]. However, the coefficient is 30 percent higher for non-EU than for EU countries. Second, the average Grubel & Lloyd index is 0.43 in the EU and 0.28 for the others. This, together with the previous point, suggests that the higher level of intra-trade in the EU is more due to the relative higher homogeneity of the economic structure in the EU than elsewhere than from some dynamic process due to the economic integration. Besides, this could be the sign of a general movement of convergence of the trade structure in OECD countries.

[24] The results are not reported here but are available upon request.

2.4 Distinguishing vertical and horizontal IIT

Fontagné & Freudenberg (1999) -henceforth FF- use the classification developed by Abd-El-Rahman (1986, 1991) as an alternative technique to the Grubel & Lloyd index. We will use this method below. It allows for a distinction between horizontal and vertical IIT. Using the terminology of FF, we will talk about two-way trade of different 'varieties' for the former and two-way trade with differences in the 'quality' for the latter. Horizontal IIT represents trade in goods of the same type that differe only by some minor characteristics. This corresponds to the differenciation of products in a monopolistic competition framework. An example could be Nike against Reebok tennis shoes. Vertical IIT corresponds to exchanges of goods presenting a significant difference of quality. It is assumed that quality is proxied by difference in prices (i.e. difference in unit-values): a good significantly more expensive than another is assumed to be of a better quality. When this difference is large enough, VIIT is somehow an intermediate case between one-way and two-way trade. An example could be Rolls-Royce against Fiat cars. FF further assume that differences in qualities are associated with differences in production functions: two economies experiencing vertical IIT should be affected by asymmetric shocks. Their main finding is that the EMU should substantially increase the share of intra-industry trade in intra-EU trade. Both types of IIT should increase, with a faster rate for HIIT.

In this part we conduct an experiment on the relation between IIT and business cycles correlations. In the previous part, we have focused on the left-hand side of the relation, by providing additional measures of business cycles 'similarity'. We now try to focus upon the measurement of the trade structure, by following FF. For this purpose, highly disaggregated data on bilateral exchanges is needed. The CHELEM database used in the previous section is not suitable for this exercise. Therefore, we turn to the Comext trade database published by Eurostat, for Belgium, France, Germany, Great-Britain and the Netherlands. The level of disaggregation is four digits[25]. This should be enough to implement this methodology, which requires data at the products level. The counterpart is that the data starts in 1995 on an annual basis, which reduces considerably the possibilities for measuring business cycles similarity. In practice, we cannot go beyond simple correlations.

[25] SH4 nomenclature. This represents roughly 1,600 products. Fontagné & Freudenberg use more highly disaggregated data. They use the NC8 nomenclature which gives detailed data for 10,000 products.

2.4.1 Methodology

Three trade types are considered: 'one-way trade', IIT where products differ by their unit value and IIT where products only differ by their variety. The methodology consists of finding the share of each type in total trade.

Let imports of country i from country j at time t for good k be $M^k_{ij,t}$. The procedure works as follows:

1. Look in the data for corresponding exports[26], that is: $M^k_{ji,t}$
2. Determine whether there is inter-industry trade or IIT for this particular observation. The procedure chooses IIT if the following condition is met[27]:
$$\frac{Min(M^k_{ji,t}, M^k_{ij,t})}{Max(M^k_{ji,t}, M^k_{ij,t})} > 0.1 \tag{2.15}$$
3. In case of IIT, determine whether it is *vertical*, i.e. if goods of type k imported and exported are of different quality, or *horizontal*, i.e. if goods are of similar quality and are simply different varieties of the same product). The latter is chosen if
$$\frac{1}{1.15} \leq \frac{UV(M^k_{ij,t})}{UV(M^k_{ji,t})} \leq 1.15$$
where $UV(M^k_{ij,t})$ represents the unit value of good k imported by i from j at time t.
4. Redo steps 1-3 for each good, holding ij and t fixed, and compute the share of trade of each type in total trade
$$\frac{\sum_{k\in K_T}(M^k_{ij,t} + M^k_{ji,t})}{M^{total}_{ij,t} + M^{total}_{ji,t}}$$
where K_T represents the set of goods classified under trade type τ [28].
5. Redo steps 1-4 for each pair ij and each t.

[26] As in the previous part, we only use data on imports.

[27] We take here the same rules as FF. However, they use a higher level of disaggregation. Therefore, there should be higher trade overlap and more IIT for our database, so that the average level of one-way trade should be lower. However, the results in terms of significance of the relation between business cycles correlations and trade should not be affected.

[28] K_T varies with t and ij. $\tau = VIIT, HIIT$ or OWT.

We now look at the relation between trade types and business cycles correlations. As in the previous part, business cycles are computed from monthly industrial production series published by the OECD. The series are transformed by the Baxter & King bandpass filter[29]. Similarity is measured by simple correlations of these filtered series, over one year. That is, the value for each year corresponds to the correlation computed over the 12 observations of that year.

Three equations are computed, each corresponding to one trade type. The following IV regressions[30] are used:

$$Co_{ij,t} = \alpha + \beta S^{\tau}_{ij,t} + u_t \tag{2.16}$$

$$S^{\tau}_{ij,t} = A \underset{q\times 1}{Z_t} + v_t \tag{2.17}$$

where $S^{\tau}_{ij,t}$ designates the share of trade type τ in total trade. Z_t is a vector of instruments. Table 2.10 describes them. We use different compositions for Z_t, either taking the instruments of the previous part (dummies for common language and common border, and log of distance between countries[31]) or using the variables of Fontagné & Freudenberg. For the latter case, we did not take all the variables used by the authors. The reason is that we need instruments, which means that they should be correlated with the independent variable of the main equation but not with the dependent one. This is why variables such as the Balassa-Bauwens normalized difference in GDP ("*Potential for external economies*") have not been used. Indeed, it is probable that a large difference in GDP would influence the correlation of business cycles. A small economy might be influenced by a bigger one –through other channels than the trade structure. Similarly, foreign direct investments from i to j might be correlated with business cycles similarity. Firms from booming economies invest abroad, which stimulates the economy of countries where this investment takes place.

Inversely, we keep some variables such as the market size (average GDP). This variable might influence trade variables, but not business cycles correlations. There is no reason to think that the German and French business cycles will be more correlated than the

[29] with cutoff periodicities at 6 and 32 quarters.

[30] Conversely to the previous part, we do not present simple OLS regressions where instruments are directly included into the equation. This specification has been tried but no coefficient was significant.

[31] First stage estimations have also been tried with the weighted distance measure published by the CEPII (www.cepii.fr). The results were almost similar and have not been reproduced here.

Belgian and the Dutch ones just because the former two economies are bigger. If there is a big difference in income per capita, one might suspect less similar business cycles, but only to the extent that the structure of production is not the same between 'rich' and 'poor' countries. Therefore, this fact would be captured by IIT variables. Consequently, the difference in income per capita should be a suitable instrument.

Variables	
HIIT	Horizontal intra-industry trade (point 3 of procedure)
VIIT	Vertical intra-industry trade
OWT	One way trade (point 2 of procedure)
Borders	Dummy variable for common borders
Language	Dummy variable for common language
Distance	Logarithm of average distance between main cities of i and j
Market size	Average GDP between i and j
Demand for differentiation	Average income per capita between i and j
Comparative advantage	Difference in income per capita between i and j

Table 2.10: Definition of instrumental variables (Z_t)

2.4.2 Results

The results for the first step equation corresponding to eq.(2.17) are presented below[32].

First stage estimations, time & group effect

	Horizontal IIT		Vertical IIT		One-way trade	
Borders	0.020	0.200 ***	0.124 ***	-0.032	0.038 ***	-0.095 ***
Language	0.074 ***	0 ??	-0.115 ***	0 ??	0.007	0 ??
Log(distance)	0.047 ***	0.118 *	0.086 ***	0.139 ***	0.014 ***	-0.060 ***
Market size		-1.50E-07 *		1.75E-07		2.16E-08
Dem. for differ.		-2.32E-05		-1.60E-05		2.87E-05 ***
Comp. adv.		-8.22E-06		-6.39E-06		-1.58E-06
R2	0.67	0.7	0.71	0.72	0.85	0.87

Table 2.11: First stage regressions on instruments

Using the instruments of FR only is insufficient since the coefficients for additional

[32] t-tests are computed using standard errors robust to heteroskedasticity.

instruments are often significant. Therefore, we will privilege specifications with all the instruments. Table 2.12 presents the regressions that include time and group effects as this was the specification the highest coefficients of determination.[33] The other models are presented in appendix A.2.2.

Second step estimations are presented below. Different estimation techniques have been tried in order to assess the sensitivity of the results. 'Between' estimations could not be computed here since they imply a transformation that reduces too much the number of observations.

IV estimations.

		Horizontal IIT	*Vertical IIT*	*One-way trade*
Simple pooled estimation:		-0.101	2.810 ***	-1.933 **
Fixed effects:	Time effect	-0.150	2.300 **	-1.887 **
	Group effect	0.051 ***	0.167 ***	-0.066 *
	Time & group effect	0.045 ***	0.140 ***	-0.051 *
Within estimation:	Simple	-0.230	2.343 **	-1.677 **
	Time effect	-0.397	2.011 *	-1.283

Dependent variable: Corr(ij,t)
Explaining variable: HIIT, VIIT or OWT

Table 2.12: Second stage regressions

The main result is that the relation is stronger with VIIT than with HIIT. The coefficients for the latter are not significant in most cases. The estimations for one-way trade are intuitive. An increase in the share of this type of trade decreases the level of business cycles correlation. This is in line with the idea that one-way trade is associated to asymetric supply shocks.

We compute also the same regressions on transformed explaining variables. They are centered around their mean and divided by their standard deviation. This allows to make comparisons between coefficients. The reason is that if an exogenous series is higher on average than another, its coefficient will mechanically be lower, even if its predicting power

[33]Some of the coefficients could not be calculated, because of the singularity of the variance-covariance of the independent variables. This is probably due to the colinearity occuring because of the reduced number of degrees of freedom. This is the case in particular when a 'group effect' or 'group and time effect' were included.

is the same. Since VIIT is clearly higher than HIIT (.584 on average against .229), it might be interesting to see what happens when variables are transformed. Results are provided in table 2.13.

IV estimations. Transformed variables

		Horizontal IIT	*Vertical IIT*	*One-way trade*
Simple pooled estimation:		-0.007	0.240 ***	-0.124 **
Fixed effects:	Time effect	-0.011	0.197 **	-0.121 **
	Group effect	0.187 **	0.209 **	-0.224 ***
	Time & group effect	0.150 **	0.205 **	-0.212 ***
Within estimation:	Simple	-0.017	0.200 **	-0.107 *
	Time effect	-0.029	0.172 *	-0.082

Table 2.13: Second stage regressions cont.

The estimations lead more or less to the same interpretation. Coefficients for HIIT are not significant, except for the ones including group effects. An interesting point is that even for such models, the coefficient of vertical IIT is higher than the corresponding one for horizontal IIT. However, the difference between coefficients associated to HIIT and VIIT is reduced compared to estimations with raw variables.

The main finding of this study is that there is a positive and significant relation between vertical IIT and business cycles correlations. This is striking at first sight in that it contradicts the idea that VIIT, being associated to differences in production functions, should entail lower business cycles correlations.

An interpretation of this fact could be the following. Even though 'two-way trade in qualities' refers to products issued from different production functions, they are produced by very similar industries, such that they could still be affected by symmetrical supply shocks. Another argument could be found on the demand side. As revealed by first stage estimations, HIIT seems more sensitive to geographic variables (distance, common borders, common language) than VIIT. An interpretation could be that different varieties of a product are quite substituable if they are of similar quality. Hence, consumers are indifferent between products and might tend to prefer buying at home, provided that geography matters. In that case, the volume of trade should be low. Under this hypothesis, a demand shock occuring at home would hardly be transmitted to a partner country having exactly the same production structure (i.e. if HIIT is the only type of trade between the two economies). Inversely, it might occur that products of the same type, but with sufficiently different qualities, would not be substituable. Consequently, consumers would be ready to buy foreign products even though geography was raising the costs of acquisition of these products. Consumers would be ready to pay more for products of a quality unavailable at home. Thus, products differentiated by their quality could increase the transmission of demand shocks.

The results presented above must be interpreted carefully. Several extensions would be needed in order to validate the conclusions. An obvious extension would be to use a more complete database -the nomenclature NC8 has more than 10,000 products- for more countries and over a longer period. The resulting increase in the number of observations would allow a more subtle measurement of business cycles synchronization.

Another extension concerns instruments. The ones selected here have less explanatory power for VIIT than for the two other trade type. Consequently, it might be useful to find other adequate instruments for this trade type. As far as the instruments do not explain VIIT well, IV coefficients might be weaker in reality than they appear.

2.5 Conclusion

This article focuses on the trade structure as an explanation of business cycles transmission. Instead of investigating the influence of trade flows, we have taken intra-industry trade as the independent variable of the equation. The reason is that trade flows alone can theoretically lead either to more coordinated cycles if trade is mainly balanced in each sector, or more idiosyncratic cycles, if trade is mainly specialized. The idea behind this study is that the structure of trade can be seen as a proxy for the output structure of a country. In other words, if the IIT index is high for a pair of countries, their output should be influenced by the same shocks and their business cycles should comove.

We have focused on two definitions of the business cycles and we have used cross-spectral densities. This tool was shown to be practical for the study of cycles transmissions. Indeed, spectral methods allow to sort out comovements of the cycles –strictly speaking, i.e. to what extent the series are dominated by the same frequencies– and synchronizations. We have looked at the relation between IIT and the measure of comovements at first and between IIT and the measure of (non)-synchronization subsequently. The first relation was expected to be positive and the second, negative.

Three points are worth noticing after this study. First, the two definitions of business cycles that we have used here often lead to different estimations and sometimes even opposite ones. For this reason, we have focused on the second definition in the interpretations, which seems to be in line with the literature. Second, we find that relations have the adequate sign and are significant most of the time. Third, however, the relation is sometimes weaker when an dummy equal to one for EU countries is included in the regressions. This last point suggests that variables other than trade or the trade structure might be of importance to explain the similarity of business cycles, for instance financial integration or the similarity of budgetary and monetary policies etc.

Chapter 3

A Common Cycle Approach of the UK/Euro zone Business Cycle Relations.

Should the UK join in?

Abstract

We use a structural model estimated by the Kalman filter in order to extract the common cycle for different groups of OECD countries. We try to evaluate to what extent the Euro zone common cycle is affected by the inclusion of the UK into this group. An important result is that adding the UK to the Euro group does not lead to a greater heterogeneity of the group as a whole. Besides, the UK business cycle is not very different from those of the Euro zone. Another point is that the influence of the UK on the 'Euro-plus-UK' common cycle is less obvious for output than for consumption, public expenditures or investment series.[1]

[1]This chapter is a modified version of Garnier (2004)

3.1 Introduction

The 2003 HM Treasury report for the assessment of the *Five Economic Tests*, put emphasis on business cycles coordination between the UK and the Euro zone. The UK cycle idiosyncrasy was underlined (Artis, 2003). However, some authors –e.g. Massmann & Mitchell (2002), Hall & Yhap (2003)– have recently pointed out that the UK cycle was getting closer to continental Europe. Apart from its political implications, the debate about UK/Euro zone business cycle relations is of importance for economists because it is directly linked to the theory of Optimal Currency Areas (henceforth OCA). Indeed, the debate about the desirability of the Euro zone goes hand in hand with this issue. The basic form of the OCA theory tells that a monetary zone is optimal if the business cycles of its members are coordinated, if there is a high mobility of factors between these members and if they trade a lot with one another. The coordination of business cycles is therefore a necessary –although not sufficient– condition for optimality and the HM treasury did consider this issue with great care. The problem is that the theory is rather vague about the degree of business cycles coordination necessary to have an optimal zone. Concerning the UK/Euro case, one way could be to make the –rather strong– assumption that the EMU is an OCA. Then, it would suffice to compare the degree of business cycles coordination within the Euro zone with that between the UK and the Euro zone, to have an idea about the desirability of the entrance of the UK in the EMU.

Many papers have addressed the issue of business cycles synchronization within the European Union. A common finding is that business cycles have become more similar with the European monetary integration process, e.g. Artis & Zhang (1997, 1999), Artis, Kontolemis & Osborn (1997)[2] . The results of Frankel & Rose (1997) go in the same way. At the same time, the UK is found to be more correlated with the US than with the other European countries. But this result appears for data starting in the 60s or 80s. Instead, the papers of Massmann & Mitchell (2002) and Hall & Yhap (2003) point out that larger business cycles co-movements between the UK and the other European countries have occurred during the past decade. More precisely, this phenomenon seems to happen after the German reunification and the European currency crisis periods.

Concerning methodology, we will use the Kalman filter and state-space modelling in

[2] An exception can be found in Inklaar & Haan (2001)

order to detect common cycles for different groups of OECD countries. This approach can be linked to the litterature on dynamic factor analysis[3].

Section 2 presents the model used and section 3 applies it to a group of OECD countries with a special attention to the UK and the Euro zone. The last part concludes.

3.2 Modelling approach

We will use in this study Kalman filtering techniques in order to extract common and idiosyncratic cycles from the series. When using *ad hoc* tools such as the Hodrick-Prescott filter, the trend will be the output of the filter, while the cycle is simply the difference between the series and this trend. The Kalman filter somehow allows a more subtle extraction of these elements. Assumptions about the behaviour of the unobserved components are made -about their law of motion- and the Kalman filter optimally extracts these components given these assumptions. This makes the interpretation of the filter output easier since a structure is put onto the model before extraction. Many different types of components other than the trend or the cycle can be extracted from the series. This structure can be built upon economic theory. See Laubach & Williams (2001) for an application to monetary policy.

We use the structural approach of Harvey (1989) and Harvey & Jaeger (1993), by decomposing the series into a so-called *local linear trend* and a stochastic cycle. This model has received important attention in recent years. Azevedo et al. (2003) propose an interesting development. They argue that leads and lags relationships are not adequately taken into account with this approach. They use a similar, but more general specification of the model, similar to that of Harvey and Trimbur (2003), which allows to produce cyclical components having the same properties as band-pass filters. Moreover, they modify the cycle in order to take phase-shifts into account as in Rünstler (2003). Cubbada (1999) extends the concept of common cycles when the series exhibit unit roots at the zero and at seasonal frequencies.

Maravall (1995) shows that two main approaches are used in the unobserved component framework : in the model based approach, classical ARIMA models are rearranged into a

[3]See, e.g. the seminal article of Forni et al. (2000). Our task here is much more modest since we use a far smaller number of dependent variables and that we restrict the number of common factors to two.

state-space form and are estimated with the Kalman filter. In the structural time series approach, each state variable has a predefined structure. He points out the flexibility of the unobserved components as a tool and shows that stochastic trends and seasonal components can adequately be estimated. He shows also that ad-hoc filtering may lead to spurious cycles.

We will use below a multivariate model with country-specific and common elements. Let y_t be a $k \times 1$ vector composed of macroeconomic series of a given group of countries, with $\mathbf{y}_t = \begin{pmatrix} y_{1,t} & \cdots & y_{k,t} \end{pmatrix}'$. The model for a given country i is assumed to be

$$y_{i,t} = \mu_{i,t} + \psi_{i,t} + \theta_i \widetilde{\mu}_t + \omega_i \widetilde{\psi}_t + \varepsilon_{i,t}, t = 1, ..., T, i = 1, ..., k \tag{3.1}$$

where $\mu_{i,t}$ and $\psi_{i,t}$ are the idiosyncratic trend and cycle, respectively, whereas $\widetilde{\mu}_t$ and $\widetilde{\psi}_t$ are the common trend and common cycle. θ_i and ω_i are coefficients specific to country i, since the common elements are unlikely to affect equally every country. $\varepsilon_{i,t}$ disturbances are assumed to be $n.i.d.$

Heuristically, the trend is supposed to be a random walk with drift and the cycle is a sum of sine and cosine functions moving at a particular periodicity. See appendices for a full description of the model. We focus on the cyclical elements $\psi_{i,t}$ and $\widetilde{\psi}_t$. The procedure is the following: coefficients are first estimated by maximum likelihood. Second, the deJong & Penzer test for structural breaks and outliers is computed, before including corresponding dummies and re-estimating the coefficients in a third step. Finally, the Kalman filter/smoother is run in order to extract unobserved cycles.

3.3 UK/Euro business cycles relations

We try to see in this empirical study whether the UK business cycle is getting closer to the one of the Euro zone. Recent papers have found results in this direction (Massman & Mitchell, 2002, Hall & Yhap, 2003). An experiment is conducted where the UK is included into the group of Euro countries[4] . We examine how the characteristics of the cycles vary

[4]By 'Euro zone', we consider in fact France, Germany, Italy, Belgium and the Netherlands, which could be seen as a 'core' group of the Eurozone. The reason is that the number of parameters to be estimated is equal to $8N+5$ (N being the number of countries), which becomes untractable with maximum likelihood estimation if the number of countries is too large. Note that these five countries represent 80 to 85% (depending on the years) of total GDP of the EMU, so that we consider this group as a proxy for the Euro zone.

when the UK is included into or excluded from this group. Another question is to see whether the UK cycle is more correlated with the Euro zone or with the US.

Following Kose et al. (2003), we look first at the share of the common part in the total cycle variance, for each country of a given group. In a second step, spectral densities are computed to see how similar individual and common cycles are. Finally, we use correlation functions in order to see how the links between the UK, the Euro zone and the US cycles have evolved in the past two decades.

3.3.1 Data

The data comes from the *OECD statistical compendium*. We use a panel composed of Belgium, France, Germany, Italy, the Netherlands, the United Kingdom, Japan and the United States for output and three of its expenditure components: consumption, public expenditures and gross fixed capital formation –we will simply refer to this variable as 'investment', below. Series are expressed in constant prices and are indexed with base 1995 Q1 = 100 and are quarterly. Whenever this was possible, the series were chosen without seasonal adjustment since this could bias the results of UC models. We follow Maravall (1995) on this point. The sample goes from 1980 Q1 to 2002 Q4.

For Germany, some intrapolation was made from the series since it only started in 1991. The series for west-Germany only was available up to 1997 for GDP and 1998 for Consumption. By taking the index of the series for west-Germany it was possible to make a 'backward interpolation' of the series for Germany. This is not realistic at first sight but is not really problematic in our case since we are more interested in the variations of the series than in their levels and that we are looking for the cyclical link among EU/EEC members. Before 1991, it is therefore the behaviour of the series of the Federal Republic of Germany that is relevant. Unfortunately, for public expenditures and investment the series for west-Germany were not available to the author, so this country has been dropped from the analysis in order not to limit the sample too much –which would have started otherwise in 1991.

3.3.2 To what extent are the cycles determined by the common factor ?

Share of the common part in the total cyclical variance

We use below the percentage shares of common cycle variance in the total cyclical variance, following Kose et al. (2003). The aim is to see how much the variability of the cycle is explained by the common part. For country i, this is given by

$$S_i = \frac{\omega_i^2 \sigma_{\tilde{\psi}}^2}{\sigma_{\psi i}^2 + \omega_i^2 \sigma_{\tilde{\psi}}^2 + 2\omega_i \mathrm{cov}(\psi_i, \tilde{\psi})} 100$$

where $\sigma_{\psi i}^2$ and $\sigma_{\tilde{\psi}}^2$ are the variances of the idiosyncratic part and of the common part of the cycle, respectively, whereas ω_i is the loading factor for country i. We also use an alternative measure that does not take the loading coefficient nor the covariance into account

$$S2_i = \frac{\sigma_{\tilde{\psi}}^2}{\sigma_{\psi i}^2 + \sigma_{\tilde{\psi}}^2} 100$$

$S2_i$ is a simplified measure that aims at correcting biases that could happen in S_i. In particular, this could be the case if a low loading factor is induced by non-synchronization between the idiosyncratic and the common cycle. Recall that this coefficient measures the contribution of the common cycle to the series at time t but does not take into account lagged comovements. The disadvantage of this measure is that its interpretation *per se* is less obvious than for S_i. However, it simplifies cross-country comparisons.

Table 3.1 compares S_i for three groups of countries, Euro, Euro-plus-UK and UK/ Japan/ US. The third group is used as a control. Table 3.2 does the same for $S2_i$.

For output, the fact of including the UK into the Euro group does not modify dramatically the value S_i. If the UK was completely disconnected from the Euro zone, there should be a decrease in the share of the variance due to the common cycle, since the group would be more heterogeneous. On the contrary, the inclusion of the UK increases S_i for two countries of the group (Belgium and Germany) out of four. At the same time, the level of S_{UK} (20.28%) is higher than that of the Netherland, Italy and Belgium, and is just below the average of the group. This suggests that the UK plus Euro cycle reflects some homogeneity. The group composed of the UK, Japan and the US has been computed as a benchmark. It reveals that the average share of the cyclical variance explained by the common cycle is around 7%, which is below the figures for the Euro and Euro-plus-UK groups.

For consumption, the values of S_i for the Euro group are much lower than those for output, suggesting that the group is less determined by a common cycle. Besides, the inclusion of the UK induces a large increase in S_i for Italy but no similar variation for the other countries. This could be interpreted as an increase in the heterogeneity of the group. Similar observations could be made for public expenditures and investment.

For the UK/Japan/US group, the values of S_i are roughly in the same range for output –between 4.8 and 9%– whereas there is much larger differences for the other macroeconomic aggregates, confirming the idea observed for the Euro and Euro-plus-UK groups that consumption, public expenditures and investment are less influenced by common cycles than output.

Percentage share of common cycle variance.

	Output			*Consumption*			*Public Expenditures*			*GFCF*		
	Euro	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*
Belgium	11.48	12.61		5.82	1.10		0.17	3.25		6.82	0.85	
France	61.85	59.29		1.68	2.96		7.43	40.83		9.85	34.49	
Germany	19.47	28.33		0.18	0.06							
Italy	20.96	16.90		0.10	37.16		32.20	69.19		9.98	25.07	
Netherlands	0.49	0.22		8.22	3.78		1.01	0.51		0.89	0.53	
UK		20.28	9.03		25.37	2.52		9.66	8.35		85.51	11.83
Japan			4.83			1.96			20.31			0.27
US			7.71			27.93			12.85			0.01
Average	22.85	22.94	7.19	3.20	11.74	10.80	10.20	24.69	13.84	6.88	29.29	4.04

Table 3.1: S_i measure of Kose et al.(2003)

A critic could be addressed to this technique in that the common cycle variance is a function of the loading factor ω_i, which measures the contribution of the common cycle to the series at time t. It might be that there is a phase lag between the cycle of the series and the common cycle, even though the series comove. This is the case in particular for the Dutch output series. One would expect the common Euro cycle to influence a lot the cycle of this country, since it is quite small compared to its partners. However, S_{NL} is well below the average of the group. This can be explained by the value of the loading factor (-0.05), reflecting the fact that the comovement of the idiosyncratic part with the common cycle is low (see table 3.3 below) and that there is a lag between the two (table 3.4).

Table 3.2 measures the share of the common cycle variance without taking the loading factor into account. Consequently, the common cycle might be oversized for this measure, but is useful for cross-country comparison. The inclusion of the UK into the Euro group does not lower the share of the common cycle, suggesting that the 'Euro plus UK' group is not more heterogeneous than the Euro group. Note that the share of the common cycle for the Netherlands output series is above the average of the Euro group.

Percentage share of common cycle variance.

	Output			*Consumption*			*Public Expenditures*			*GFCF*		
	Euro	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*
Belgium	22.51	23.98		18.76	8.99		11.72	19.79		36.72	36.21	
France	89.94	86.01		45.43	45.37		64.74	86.14		92.82	95.49	
Germany	36.80	47.59		43.71	36.66							
Italy	47.20	38.99		7.94	69.70		88.12	97.71		54.31	69.28	
Netherlands	63.29	65.40		19.03	16.90		27.05	41.90		9.48	10.08	
UK		71.69	36.54		43.12	57.70		28.91	30.89		98.80	56.04
Japan			15.96			9.60			63.23			4.24
US			75.91			82.54			40.28			4.63
Average	51.95	55.61	42.80	26.97	36.79	49.94	47.91	54.89	44.80	48.33	61.97	21.64

Table 3.2: Second measure $S2_i = 100\sigma^2_{\hat{c}}/(\sigma^2_{\hat{v}i} + \sigma^2_{\hat{c}})$

If one looks at the simple ratio of common variance over idiosyncratic variance (not reproduced here), the value for the Netherlands is in the same range as the other countries. This shows the importance of properly taking phase lags into account (section 3.3.3).

Spectral densities

Another useful information is given by spectral densities. By comparing the spectra of the idiosyncratic and common part of the cycles, one might get some insight about the homogeneity of the group, depending on whether the cycles are dominated by the same frequencies. By contrast, it could happen that the group is influenced by one country only. In that case, the common cycle would exhibit a peak at the frequency at which this particular country oscillates.

The plots for the Euro-plus-UK group only have been represented below. See appendix A.3.6 for the Euro group. The spectral densities were computed using a Parzen window with a truncation parameter of 10.[5]

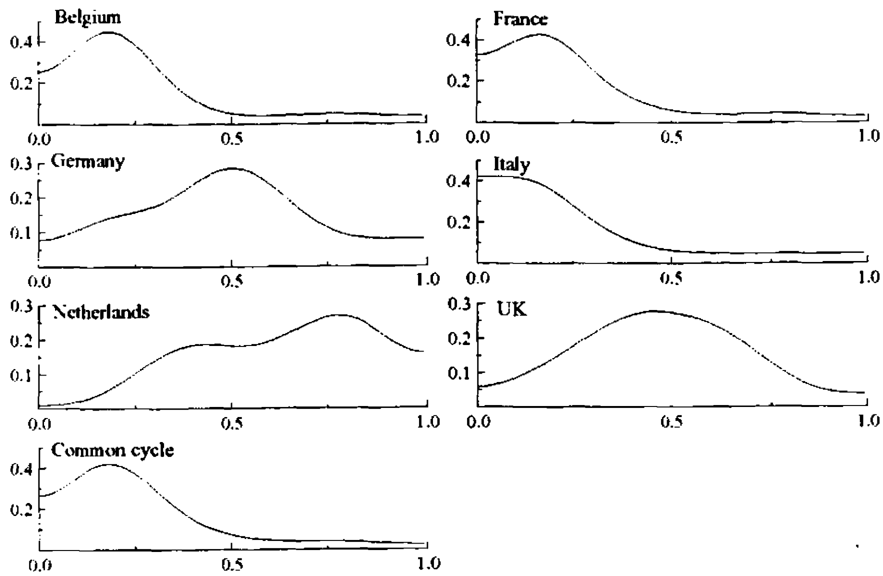

Figure 3.1: Spectral densities for 'Euro plus UK' group – Output

For output, the spectrum of the common cycle is dominated by frequencies around 0.2.[6] This shape is pretty similar to that of Belgium, France and to a lesser extent, Italy. The shape of the spectral densities is quite different for the UK, but this is the case for Germany, and for the Netherlands as well. For consumption, the bulk of the common cycle is concentrated around the same frequencies as for output (around 0.2, i.e. a periodicity of 5 quarters) and this is similar to Belgium and Italy –and maybe Germany, although the peak is at a slightly higher frequency. Another, smaller peak can be seen at higher frequencies, suggesting the influence of other countries: France, the Netherlands and the UK. The third macroeconomic aggregate is Public expenditures. Once again the spectral shape is quite similar to that of Belgium, France and Italy. UK and the Netherlands show a different behaviour. The similarity between the UK and the Euro group common cycle is higher for investment where the two shapes look quite similar. Recall that Germany

[5]i.e. the periodogram was computed up to 10 lags/leads

[6]For convenience, frequencies have been scaled to lie between 0 and 1, i.e. the frequency is simply the inverse of periodicity. One might also find the representation between 0 and π.

has been dropped from public expenditures and investment databases.

In short, the Euro group common cycle spectral density exhibits a similar pattern as densities for Belgium, France and Italy. The UK has quite different shapes but the dissimilarity with the common cycle is not greater than for the Netherlands.

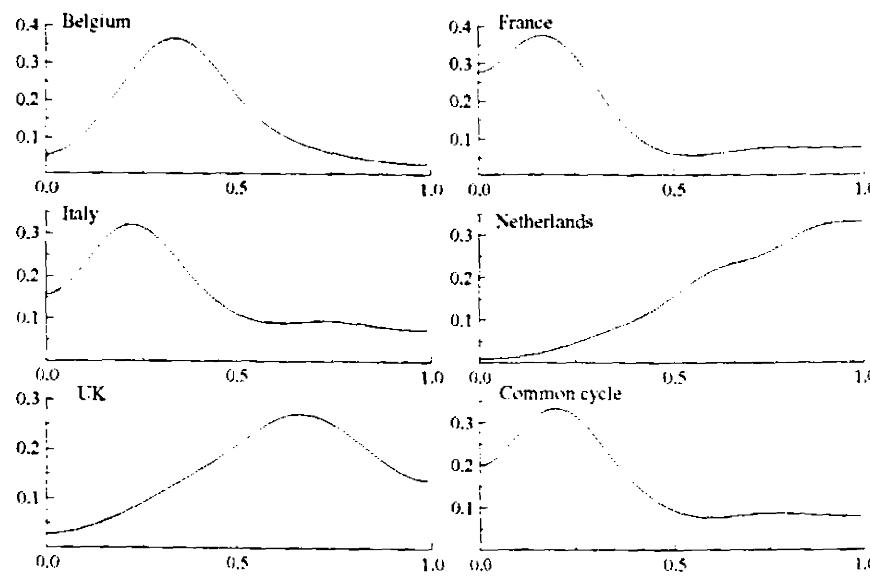

Figure 3.2: Spectral densities for 'Euro plus UK' group - Consumption

For the Japan/UK/US group –see appendix A.3.6 - the spectral density common cycle is very similar to that of the US for output, consumption and public expenditures, suggesting the dominating role of this country on the two others. For the two last variables, the influence of Japan and the UK on the common cycle is also visible. For investment, there is a surprising difference between the US and the common cycle.

3.3.3 Correlation/synchronization of the cycles

In this paragraph we use lagged *maximum* correlations in order to see how much the series comove and how much they are synchronised. It is clear that simple correlations are not sufficient from this point of view since they do not capture the lagged movements of the series. Using the above procedure, we estimate the cyclical part for each individual series and for several groups of countries. For each pair of series, correlations are computed at

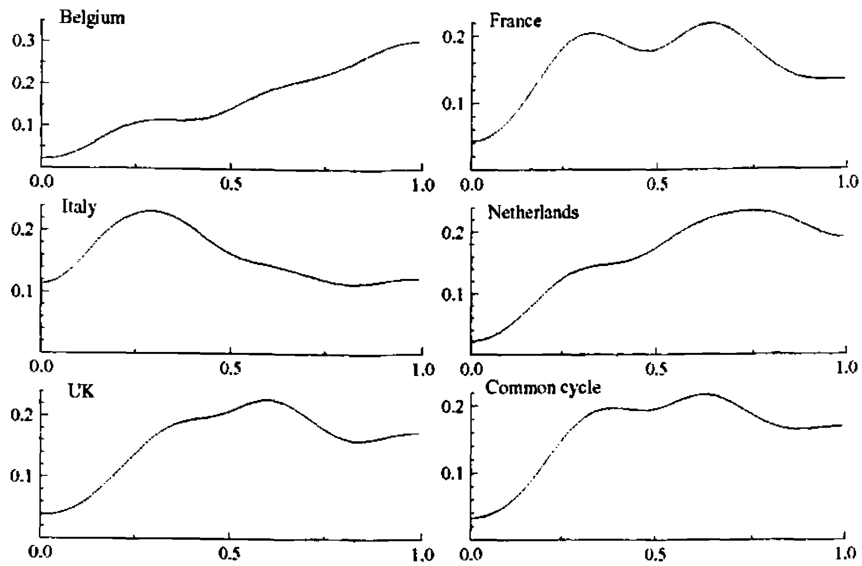

Figure 3.3: Spectral densities for 'Euro plus UK' group – Public expenditures

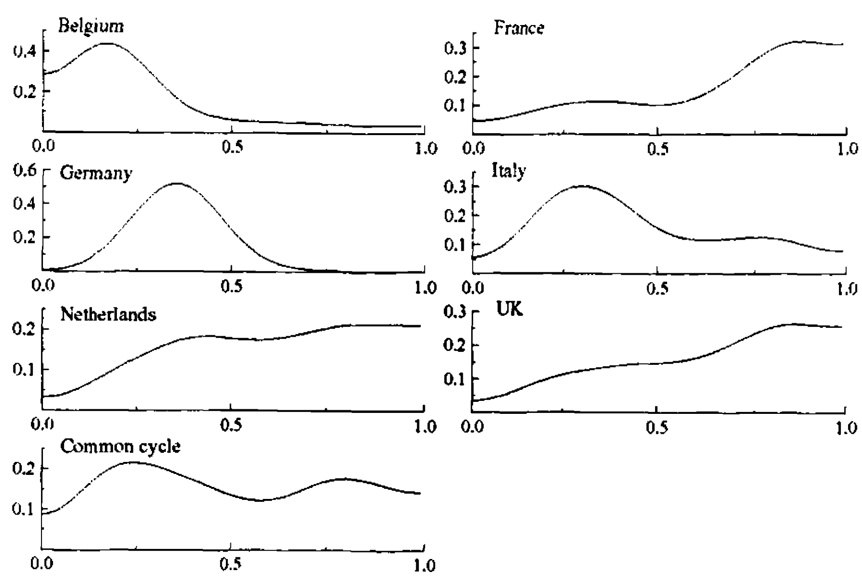

Figure 3.4: Spectral densities for 'Euro plus UK' group – Investment

different lags[7], and the maximum correlation is stored, together with the corresponding lag ($maxC_{xy}(\tau)$and $argmaxC_{xy}(\tau)$, respectively). The information contained in these two measures gives some insight about the way series comove and about their synchronization.

Correlation within groups

We examine here how the individual and common cycles are correlated within each group -Euro, Euro-plus-UK and Japan/UK/US. This will complement the information provided by variance shares. Table 3.3 below measures the maximum correlations between individual cycles ($\psi_{i,t} + \omega_i\widetilde{\psi}_t$) and the common part of the cycles ($\widetilde{\psi}_t$). See eq.(3.1). It shall be noted that $\widetilde{\psi}_t$ is a part of the two variables. Therefore, instead of focusing on whether they are correlated or not, we will concentrate on the comparison of the correlations between groups. We look at the effect of including the UK into the Euro group and we compare the Euro and the Japan/UK/US groups.

Maximum correlations between univariate and common cycles

	Output			*Consumption*			*Public Expenditures*			*Investment*		
	Euro	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*
Belgium	0.63	0.62		0.34	0.51		0.36	0.16		0.51	0.25	
France	0.79	0.79		0.23	0.19		0.40	0.32		0.20	0.50	
Germany	0.34	0.34		0.14	0.16							
Italy	0.56	0.54		0.34	0.57		0.31	0.66		0.67	0.23	
Netherlands	0.20	0.22		0.43	0.26		0.20	0.13		0.20	0.19	
UK		0.42	0.51		0.60	0.30		0.35	0.51		0.39	0.57
Japan			0.38			0.20			0.72			0.72
US			0.71			0.95			0.65			0.24
		(0.5)			(0.34)			(0.32)			(0.29)	
Average	0.50	0.49	0.53	0.30	0.38	0.48	0.32	0.32	0.63	0.39	0.31	0.51

nb: the average for Euro group only is in brackets

Table 3.3: Maximum correlations between individual cycles and the common cycle of their group

Whole Sample For output, the average correlation is higher for the Japan/UK/US than for the Euro group (0.53 against 0.50), but the figures are in the same range. Besides, the three highest correlations (Belgium, France and Italy) are greater on average than that of Japan/UK/US. Turning now to the inclusion of the UK, we see that it does not modify

[7] For two series x_t and y_t we compute the correlation function $C_{xy}(\tau) = S_{x(t)y(t+\tau)}/(S_{x(t)}S_{y(t+\tau)})^{1/2}$, $t = 1, \ldots, T$, $\tau = -n, \ldots, n$ where S represents empirical second order moments. Here we have set $n = 6$.

the figures greatly. In addition, the correlation of the UK cycle with the common cycle is higher than that of Germany and the Netherlands. This suggests a similar observation as in the previous paragraph. The UK business cycle exhibits some homogeneity with the Euro group, since its inclusion into the group does not modify much the structure of the relations of the individual cycles.

The inclusion of the UK modifies more the levels of the correlation for consumption than for output. However, correlation increases for three countries of the group, namely France, Germany and Italy. The UK cycle is more correlated with the common cycle than the other countries, suggesting that it has somehow 'attracted' the common cycle, thereby indicating an effect of the UK business cycle on the common cycle. However, the fact that the average correlation increases for the Euro group tends to show that including the UK does not entail more heterogeneity for this group. Similar remarks can be done for public expenditures and physical investment: the UK cycle is as highly correlated with the common cycle as the average of the Euro group, but its inclusion leads to a modification of the correlations structure of this group.

An important fact to be noticed is that the average *within* correlation is higher for the Japan/UK/US group than for the Euro or Euro-plus-UK groups. An interpretation, in line with Kose et al. (2003), could be that the common component at the world level plays a larger role than at the continental level.

The average lags are presented below. They measure the lag at which the correlation between the first and second series is highest. A negative value implies that the individual country leads the common cycle. For output, all the countries of the Euro group are in phase, at the exception of the Netherlands, and the inclusion of the UK into the Euro group does not affect this synchronicity. However, this is not the case for the components of output. The Euro-plus-UK group has a different lag structure than the Euro group, revealing once again how the UK modifies the common cycle for these variables. For instance, Belgium, France and Italy lead the common cycle for the consumption of the Euro group, whereas they are more or less in phase with the Euro-plus-UK group. Inversely, the Netherlands are in phase with the common cycle in the first case, and lead it in the second case. A surprising result is that the German cycle lags behind the common cycle of the Euro zone for this variable, which seems counter-intuitive.

Note that the Japan/UK/US group is more synchronized. For all variables, the three countries seem to be in phase. The only minor exceptions being consumption where Japan leads by one quarter and investment where the US leads by 3 quarters.

Average lags between univariate and common cycles

	Output			Consumption			Public Expenditures			Investment		
	Euro	Euro-UK	UK-Jp-US	Euro	Euro-UK	UK-Jp-US	Euro	Euro-UK	UK-Jp-US	Euro	Euro-UK	UK-Jp-US
Belgium	0	0		-4	0		6	0		0	5	
France	0	0		-2	1		0	4		2	0	
Germany	0	0		4	-2							
Italy	0	0		-4	0		-4	0		0	-1	
Netherlands	-1	-1		0	-3		-2	-3		-2	0	
UK		0	0		0	0		0	0		0	0
Japan			0			-1			0			0
US			0			0			0			-3

Table 3.4: Average lags between individual cycles and the common cycle of their group

Rolling correlations We use rolling correlations in order to see the evolution of the comovements and the synchronization. A window equal to 40 observations (10 years) was selected[8]. We focus on maximum correlations as explained above. Here, the evolution of the correlation of the UK cycle with the common part of the group is considered. The dates indicate the centre of the rolling window –i.e. the first date, 1985q1 indicates a correlations on the sample 1980q1-1990q1.

Four different patterns can be observed from figure 3.5. For output and public expenditures, the UK is more correlated with the Japan/UK/US common cycle than with the one of the Euro group. For the former series the difference disappears in the last years (from 1996 to 1998, i.e. over a period comprised between 1991 and 2003). For the latter series, the correlation of the UK cycle with both common cycles grows but is higher with the Japan/UK/US group. Inversely, the correlation with the Euro group common cycle is higher for consumption, although it tends to decrease. The difference between the two measures diminishes in the most recent years. Overall, it seems that the difference between the cyclical relations of the UK with the two groups tends to decrease for three variables

[8]The size of the window is constrained below by the number of lags taken into account in the correlation function. Indeed, the number of observations taken in the rolling correlation is equal to $n - \tau$, where τ is the number of lags in the lagged correlation. At the same time, n is constrained above by the sample size.

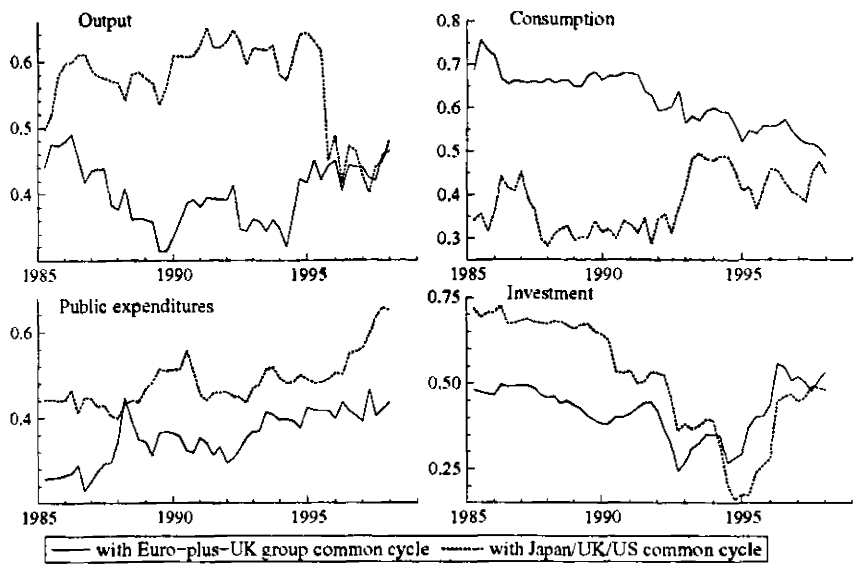

Figure 3.5: 'UK in': Rolling maximum correlation between the UK cycle and Euro-plus-UK or Japan/UK/US common cycles.

out of four. The exception being public expenditures, for which both correlations tend to increase.

We now take the problem from another perspective and look at the correlations for the Euro group. How are they affected by the inclusion of the UK? Plots of the correlations between the idiosyncratic and the common cycle for the different countries of the Euro group are displayed in appendix A.3.2. They suggest that the effect of this inclusion is only marginal for output. In fact, the UK cycle is not modified by its inclusion into the Euro group nor is the Euro common cycle affected by the inclusion of the UK. This leads to two opposite interpretations. 1/ The UK cycle is either well coordinated with the Euro zone common business cycle such that it only marginally modifies it. 2/ Or the UK cycle is orthogonal to the common cycle of this group, such that they do not share any common cycle. The results obtained for variance shares (paragraph 3.2.1) and for spectral densities (3.2.2) would be more coherent with the first interpretation. However, this issue remains open.

On the contrary, the effect is quite important for consumption, public expenditures and investment. However, there is no systematic decrease in the correlation associated with the inclusion of the UK.

Correlations between cycles for different groups

We compare now the cycles across different groups or between groups and univariate series. In the latter case, we compare the common cycle of the group - as estimated in equations (A.21) and (A.22) in the appendix- and the cycle obtained from the system (A.16)-(A.17). This will be the case for, e.g. the Euro zone and the UK.

Whole sample Table 3.5 shows contemporaneous correlations between cycles[9] –common or individual ones. The correlation Euro/US should increase when the UK is included in the former group. Surprisingly, this correlation increases only slightly for output and decreases for the components of output. Another thing is that the correlation with the group *Ja/UK/US* should be much higher for Euro-plus-UK than for Euro alone, since the UK is present on both sides. However, the increase is quite small for all variables. An explanation could be that the UK only accounts for a small share of the Japan-UK-US and Euro-plus-UK groups common cycles. This is plausible for output, since we have seen above that the common cycle of the Euro zone was not much affected by the UK, but it is a bit more surprising for other variables. Anyway, these remarks seem to weaken the view according to which the UK is closer to the US cycle than to the Euro one.

Table 3.6 is dedicated to the analysis of synchronization between cycles. The US cycle is in phase with the Euro cycle for output but leads it for investment. Note that approximately the same structure is visible for row 1 ('*Ja/UK/US*') and for row 3 (US), which shows that the common cycle of the *Ja/UK/US* group is pretty similar to the US cycle. Concerning the effect of including the UK, it does not modify the structure for output and modifies it slightly for investment. At the opposite, the time concordance between the Euro group and the US cycle is quite different from the relation between the *Euro-plus-UK* and the US cycle for consumption and public expenditures. Therefore, the UK modifies the Euro group common cycles substantially.

[9] The correlations where the UK is present on both sides – *Ja/UK/US* with *Euro+UK* and *UK* with *Euro+UK* – are presented for control.

Maximum Correlations between univariate or common cycles

	Output		Consumption		Public expenditures		Investment	
	Euro	Euro+UK	Euro	Euro+UK	Euro	Euro+UK	Euro	Euro+UK
Ja/UK/US	0.20	0.21	0.07	0.10	0.26	0.27	0.18	0.19
UK	0.28	0.42	0.16	0.60	0.22	0.35	0.19	0.39
US	0.15	0.19	0.13	0.10	0.32	0.24	0.18	0.13
Jap.	0.14	0.13	0.16	0.20	0.14	0.20	0.09	0.19

Table 3.5: Maximum correlation between cycles - Euro vs. UK/US/Japan

Average lags between univariate or common cycles

	Output		Cons.		Public exp.		Invest.	
	Euro	Euro+UK	Euro	Euro+UK	Euro	Euro+UK	Euro	Euro+UK
Ja/UK/US	-2	0	-3	1	6	-6	-4	-6
UK	0	0	3	0	6	0	1	0
US	0	0	-3	1	6	-6	-4	-6
Jap.	-1	-1	4	-1	-1	4	-2	-3

nb: a negative value means that the cycle in row leads the one in column

Table 3.6: Average lags between cycles - Relations Euro vs. UK/US/Japan

Maximum Correlations - UK/partners

	Output	Cons.	Public exp.	Invest.
Euro	0.28	0.16	0.22	0.19
US	0.25	0.19	0.24	0.27
Jap.	0.10	0.23	0.21	0.11

Table 3.7: Maximum correlations with the UK individual cycle

Average lags - UK/partners

	Output	*Cons.*	*Public exp.*	*Invest.*
Euro	0	-3	-6	-1
US	1	1	-6	-3
Jap.	-6	-1	3	-6

nb: a negative value means that the cycle in row leads the UK

Table 3.8: Average lags with the UK individual cycle

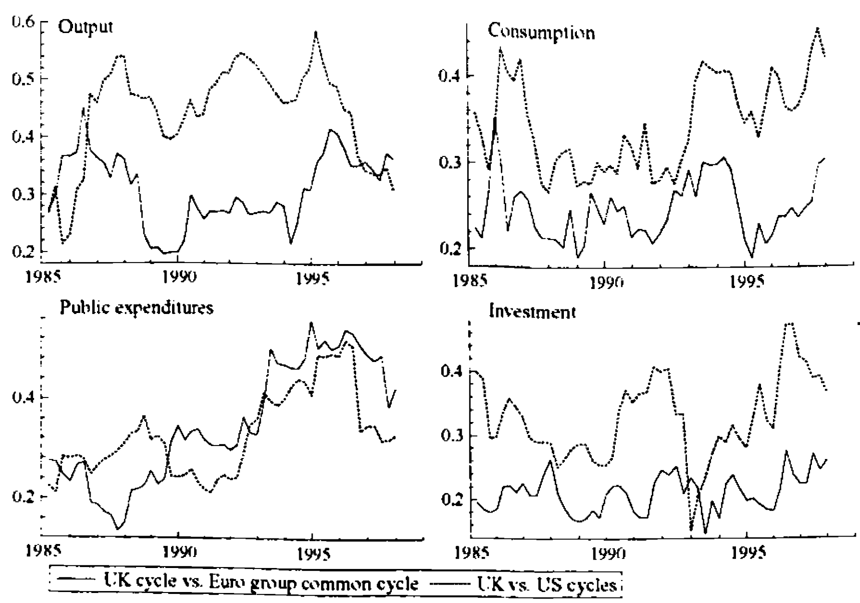

Figure 3.6: 'UK out' – Maximum correlations between the UK individual cycle and the Euro common cycle or the US cycle

Rolling correlations Figure 3.6 shows the rolling correlations between the UK and the Euro group common cycle and between the UK and the US cycles. For output, the average correlation is higher with the US than with the Euro group, confirming previous findings in the literature (e.g. Artis & Zhang, 1997). At the exception of public expenditures, the correlation for other variables is higher for UK/US than for UK/Euro zone. A noticeable feature is that the correlation with the Euro group tends to increase in the last part of the sample, in particular after 1995 (i.e. correlations calculated over 1990-2000 and after) for output, consumption and investment. This is in line with Massmann & Mitchell (2002) and Hall & Yhap (2003).

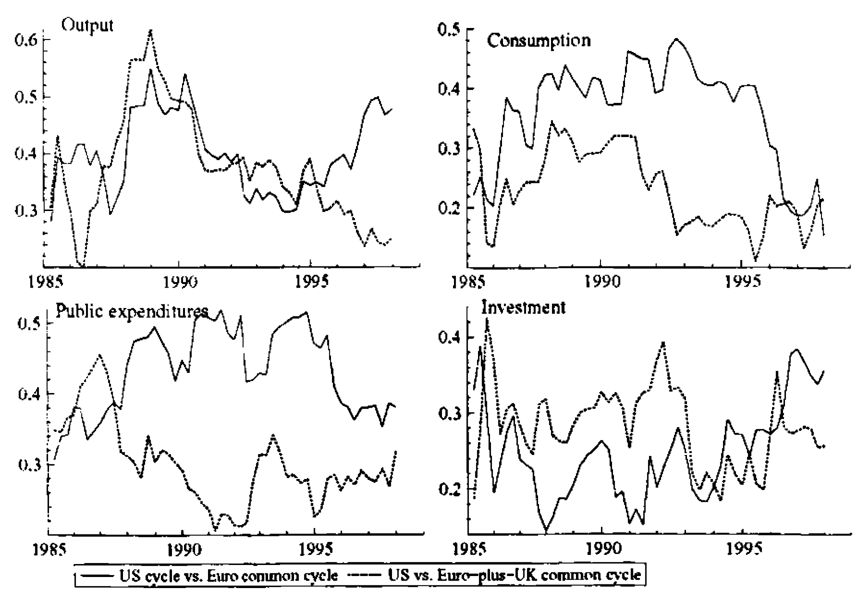

Figure 3.7: Maximum correlation of the US cycle with the Euro or Euro-plus-UK common cycle

If the UK is more correlated with the US than with the Euro group, one should normally find that including the UK into the Euro group leads to a higher correlation with the US. However, it is difficult to draw such a conclusion from figure 3.7, apart from investment before 1993. This suggests that including the UK into the Euro group does lead to a modification of the common cycle but not necessarily in a way that makes it more correlated with the US cycle. Strikingly, there is a clear divergence between the US/Euro and the US/Euro-plus-UK output correlation towards the end of the sample. Indeed, after 1995, the UK seems to attract the Euro cycle towards more idiosyncrasy *vis-à-vis* the US. This is a surprising reinforcement of the results of Massmann & Mitchell (2002) and Hall & Yhap (2003).

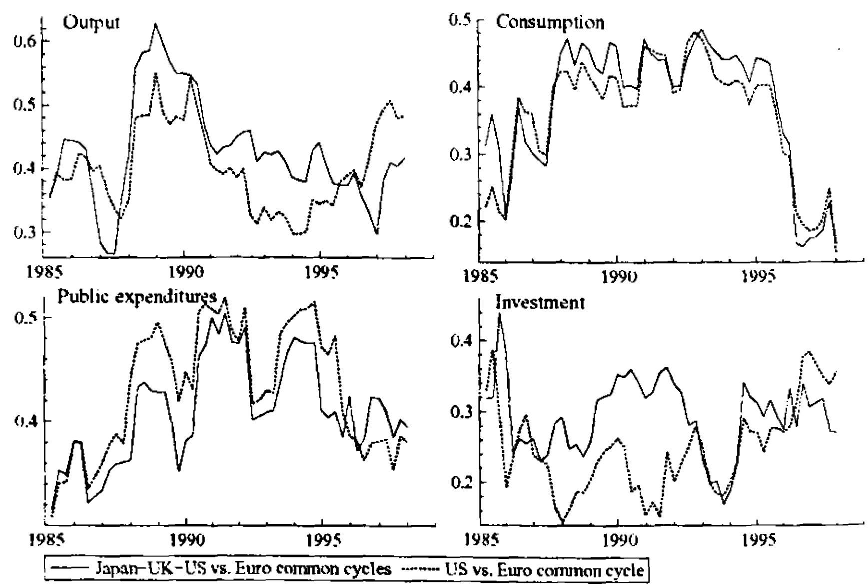

Figure 3.8: Maximum correlation of the Euro-group with the 'rest of the world' common cycles

3.4 Conclusion

We have used in this study a structural time series estimated by the Kalman filter in order to extract idiosyncratic and common cycles for different groups of OECD countries. A particular attention was given to the relation between the UK and the Euro zone. Indeed, there is an ongoing debate about the business cycle links between the UK and continental Europe.

The results for GDP series are partly in line with the existing literature. On the one hand, it seems that the UK cycle is closer to the US cycle than to the Euro area one, which is in line with the findings of e.g. Artis et al. (1997). On the other hand, the correlation with the Euro common cycle has tended to increase since the beginning of the past decade as is the case for Massmann & Mitchell (2002) and Hall & Yhap (2003). At the same time, the UK output cycle does not modify much the common cycle of the Euro zone when it is incorporated into this group.

The results for the other macroeconomic series used here –consumption, public expenditures and physical investments– diverge from those of GDP in that the inclusion of the UK into the Euro group modifies substantially the common cycle of the group. Besides, the share of the common cycle variance into the total cyclical variance is not lower for the UK than for the members of EMU, which tends to show that the UK cycle is not less determined by the common cycle than the Euro countries. Another argument is that the inclusion of the UK into the Euro-group does not lower the average correlations of the Euro countries with the common cycle. This indicates that the inclusion of the UK does not make the group more heterogenous. The spectral density of the UK cycle does not exhibit major differences in its shape from the common cycle, except for public expenditures. In any case, it does not depart more from the common cycle than the Dutch cycle or -to a lesser extent given data availability- to the German one. These results depart from existing literature.

For synchronization, Euro countries GDP cycles are in phase with the common cycle on average, except the Netherlands. Adding the UK to the Euro group does not modify this fact. However, the addition of the UK alters synchronization between Euro countries idiosyncratic cycles and the common cycle. At the same time, the UK is synchronised with the common cycle for all the variables. This suggests the existence of connections between the UK and the Euro zone business cycles.

The general result of this work is that the UK business cycle is not highly different from the Euro zone cycles. Moreover, adding the UK to the Euro group does not lead to a greater heterogeneity of the group as a whole. A second important point is that the 'UK effect' is weaker for output series than for consumption, public expenditures or investment series. This suggests the importance of taking the components of output into account when looking at international business cycles. One might also conclude that business cycles links between the UK and continental Europe have been underestimated in the literature since much emphasis was put onto output series.

Business cycle coordination can be seen as a necessary condition for having an OCA. In that case –and provided that the Euro zone is an OCA itself– the policy implication of this paper is that it is hardly possible to reject the hypothesis according to which the 'Euro-plus-UK' zone would be an OCA. In other words, under the assumption that the Euro zone is optimal, one cannot tell *a priori* that the 'business cycles condition' is not fulfilled for the Euro-plus-UK zone. Of course, postulating that the Euro zone is an OCA is quite strong. But one could make a weaker statement and see the Euro area as a 'plausible' monetary zone. The results of this paper would then suggest that the Euro-plus-UK zone might be a plausible monetary zone as well.

Chapter 4

The Natural Rate of Interest and the Output Gap in the Euro Area: a Joint Estimation

Abstract

The notion of a natural real rate of interest, due to Wicksell (1936), is widely used in current central bank research. The idea is that there exists a level at which the real short term interest rate would be compatible with output at its potential level and stationary inflation. Such a concept is of primary concern for monetary policy because it aims at providing a benchmark for the monetary policy stance. This paper applies the model of Laubach and Williams (2003) to estimate the natural real interest rate of the euro area economy over the past 30 years. The Kalman filter is run on a small-scale macroeconomic model, where inflation, the output gap and the real interest rate are related via an IS curve and a Phillips curve. The estimates suggest that the natural real interest rate in the euro area was, on average, higher than the actual real interest rate in the 1970s, while it was lower compared to the actual rate most of the time in the 1980s and 1990s. The results are broadly similar to those obtained by Laubach and Williams (2003) on US data.[1]

[1]This chapter is a modified version of an article prepared in collaboration with Bjorn-Roger Wilhelmsen (Bank of Norway and formerly ECB, Monetary Policy Stance division)

4.1 Introduction

In the long run, economists assume that nominal interest rates will tend towards some equilibrium, or 'natural', real rate of interest plus an adjustment for expected long run inflation. The natural rate of interest is a central concept in the monetary policy literature since it provides policymakers with a benchmark for monetary policy. In theory, it is an important indicator of the policy stance since rates above (below) the natural rate are expected to lower (raise) inflation.

From an empirical point of view, the 'natural' real rate of interest is unobservable and has to be estimated. The estimation of the natural real interest rate is not straightforward and is associated with a very high degree of uncertainty. In practice, therefore, policymakers cannot rely exclusively on the real interest rate gap, defined as the difference between the real short term interest rate and estimates of the natural real interest. Rather, a comprehensive approach using a wide set of information is required. This notwithstanding, central bank economists have increasingly devoted attention to developing estimation strategies for the natural real interest rate. The various methods used range from calculating the average actual real interest rate over a long period to building dynamic structural general equilibrium (SDGE) models subject to nominal rigidities.

A recent contribution to the literature on how to empirically approach the concept of the natural rate is a paper by Laubach and Williams (2003, henceforth LW) with an application on data for the United States. They suggest to estimate the natural real rate of interest and the output gap simultaneously, using a small-scale macroeconomic model and Kalman filtering techniques. In this model, the natural real rate of interest is related to output gap. Thus, the estimate of the natural real interest rate is time-varying and related to long-term developments in the real characteristics of the economy, consistent with economic theory. This method has become popular since it strikes a compromise between the theoretically coherent SDGE approach and ad-hoc statistical approaches, as emphasized by Larsen and McKeown (2002).

In this paper we employ the technique of LW for the euro area. The present work differs from others[2] in that we use a relatively long (synthetic) dataset starting in the

[2]Several papers have used this framework on European data: Sevillano and Simon (2004) for Germany, Larsen and McKeown (2002) for United Kingdom and Crespo-Cuaresma et al (2003) and Mésonnier and Renne (2004) for the euro area.

early 1960s. Besides, we estimate the natural real interest rate in Germany and the US for comparison. Such comparisons are interesting because, prior to 1999, monetary policies in Europe were considerably influenced by the Bundesbank policy. At the same time, the US is often considered as a useful proxy for global influences that affect monetary policy worldwide. Third, we apply simple statistical tests to investigate the leading indicator properties of the estimated real interest rate gap (the difference between the actual and the natural real interest rate) on inflation and economic activity.

Our baseline results suggest that the natural real interest rate has declined over the past 40 years, from around 4% in the 1960s to slightly less than 2% in 2004. At the same time, the fluctuations have been relatively low. The latter contrasts to some extent with the estimates reported in Cuaresma et al (2003) and Mésonnier and Renne (2004) on a dataset starting in 1991 and 1979 respectively[3]. This implies that our estimate of the real interest rate gap is relatively persistent over longer periods, with low short-term fluctuations. As regards the output gap, our estimates suggest that it was positive, *on average*, in the 1970s when inflation was high. Likewise, in the 1980s and 1990s, when the monetary authorities in most European countries run disinflationary policies, the output gap was negative, *on average*. Yet, the length of the business cycle's booms and busts are in line with the consensus view in the business cycle literature. Finally, it should be borne in mind that the general caveats associated with interpreting estimates of the natural rate and the output gap, which are highly uncertain, also applies for this paper.

The paper is structured as follows. In section two we take a look at the literature. Sections three discusses some facts related to the real interest rate from 1960 and presents the modeling approach. Section four displays the estimation results and briefly discusses the leading indicator properties of the estimated real rate gap. Section five concludes.

4.2 A review of the literature: short vs. long-run perspectives

The concept of a ***natural rate of interest*** was first introduced by Knut Wicksell in the late 19^{th} century (1898, with 1936 translation). Today the concept knows a revival of interest following Woodford's seminal book, *Interest and Prices*. The natural rate is integrated

[3]These papers report a peak in the natural real interest rate of around 6-7% in the early 1990s.

into a neo-keynesian framework. It is not so much the level of interest rates that should be borne in mind by central bankers, than its deviation from an equilibrium or 'natural' rate. In this perspective, the IS and Philips curves are considered in terms of deviations from long run paths. In other words, the former relates the output gap (one could also talk about business cycle) to the real interest rate gap, and the latter relates inflation deviations to output gap. This approach is exactly the one of LW.

The available estimates of the historical developments in the euro area natural real interest rate differ considerably from one author to the other. We briefly address these differences below and classify estimates of the natural real rate, taking as criterium the time horizon at which they should be interpreted.

Fluctuations in the real interest rate may be decomposed into two different components: a natural real rate and a real rate gap (Woodford, 2003; Neiss and Nelson, 2003; Cour-Thimann et al, 2004). The natural real rate is the real interest rate that would prevail in theory under perfectly flexible prices. It is related to structural factors. The real interest rate gap is related to the business cycle and reflects the existence of nominal rigidities in the economy.

Some papers find that most of the fluctuations in the real interest rate should be attributed to fluctuations in the real interest rate gap rather than the natural real interest rate. This group of papers, which includes Giammarioli and Valla (2003), Neiss and Nelson (2003), Sevilliano and Simon (2004) and LW, associate the fluctuations in the natural real interest rate with the evolution of real fundamentals such as the determinants of GDP growth and preferences. These variables are typically stable in the short to medium term, but may display some variation in the longer run. Consequently, the natural real rate is also relatively stable in the short run, and the natural real interest rate in these papers should be considered in a *long-run* perspective. It refers indeed to the level expected to prevail in, say, the next five to ten years, after any existing business cycle 'booms' and 'busts' underway have played out. Note however, that the estimated natural real rate of Mésonnier and Renne (2004) is much more volatile than that of LW or Sevilliano and Simon (2004).

On the contrary, other papers conclude that fluctuations in the natural real interest rate explain most of the variations in the real interest rate (Basdevant et al., 2004; Cuaresma et al, 2003; Cour-Thimann et al, 2004; Larsen and McKeown, 2002). The papers consistent with this view typically make use of the Kalman filter or other filtering techniques to split the actual real rate into a trend (the natural real rate) and a cyclical component (the real rate gap). However, the models they use do not necessarily contain judgement about the determinants of the natural rate. Rather, the approach they take is closer to a pure statistical measure. Consequently, variations in the natural rate are more pronounced, because the natural rate tends to follow more closely the medium term fluctuations in the actual real rate. The interpretation of the natural real interest rate is therefore likely to be more relevant in a 'shorter' time perspective in that it refers to a neutral monetary policy stance in a situation where the economy has not necessarily settled at its long run levels.

4.3 Estimating the natural rate of interest

4.3.1 Some facts

The data covers a total of 161 quarterly observations from 1963q1 to 2004q1 for the euro area, Germany and the US. The dataset consists of short-term interest rates, inflation and gross domestic products for the three economies. The computation of real interest rates is subject to several practical and conceptual difficulties. Ideally, an estimate of the real interest rate should be obtained by subtracting ex ante inflation expectations from nominal interest rate. However, the lack of good data for inflation expectations forces us to take a more straightforward approach. In this paper, the real interest rates are calculated from the three-month money market rates and annual consumer price inflation rates[4]. Hence, deviations may exist between expected and current inflation. While we believe that this problem is less severe when assessing developments over longer horizons, the deviations may be stronger in periods with unanticipated inflation, notably in the 1970s.

It should be borne in mind that monetary policy regimes differed significantly over time and across countries. Moreover, in many euro area countries, specifically in the 1960s and early 1970s, other instruments than interest rates were important in the conduct of monetary policy. In particular, capital controls prevailed in many euro area countries. Furthermore, inflation, economic growth and interest rates were very volatile in some

[4]For the euro area, national levels for interest rates and consumer prices have been aggregated prior to 1999 using GDP and consumer spending weights respectively at PPP exchange rates, see ECB (2003).

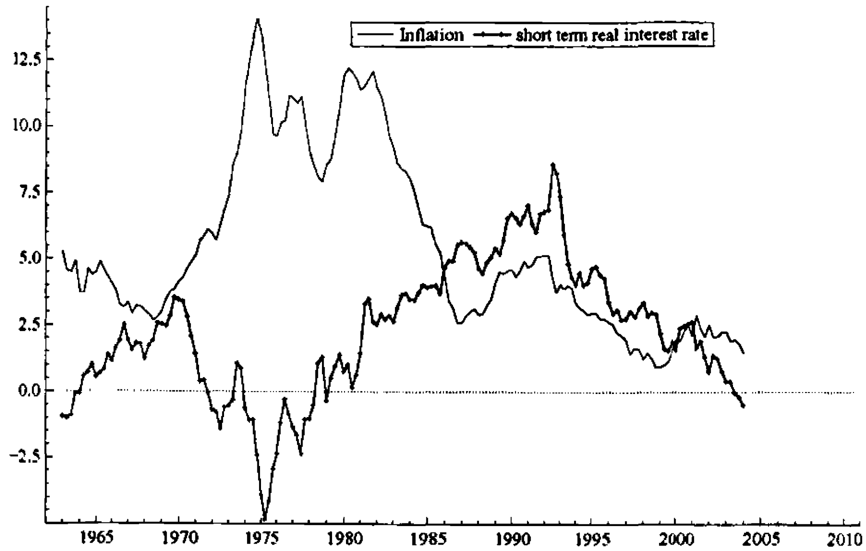

Figure 4.1: Euro area real interest rate and inflation

countries of the euro area. Finally, the euro area real interest rate has been significantly influenced by other factors than monetary policy, such as tensions within the Exchange Rate Mechanism (ERM) at the turn of the 1990s (Cour-Thimann et al, 2004).

The real interest rates fell dramatically in the 1970s when overheated economies and rising oil prices pushed up inflation at a level that could not be offset by the nominal interest rates. See figure 4.1. Following the trough in the mid-1970s, as European monetary authorities gradually put more emphasis on disinflationary policies, the real interest rate increased slowly over a period of more than 15 years. After peaking in the early 1990s, the real rate declined gradually again, influenced by the authorities' achievment of more favourable inflation developements.

Similar to the euro area aggregate, the real interest rate in the United States was also lower than its average in the 1970s and higher in the 1980s (table 4.2). However, the persistence in the data seems less pronounced, as indicated by the quick rise in the real rate at the turn of the 1980s. In Germany the real interest rate has been more stable around its long-term average, reflecting the achievment of lower and more stable inflation over the whole sample.

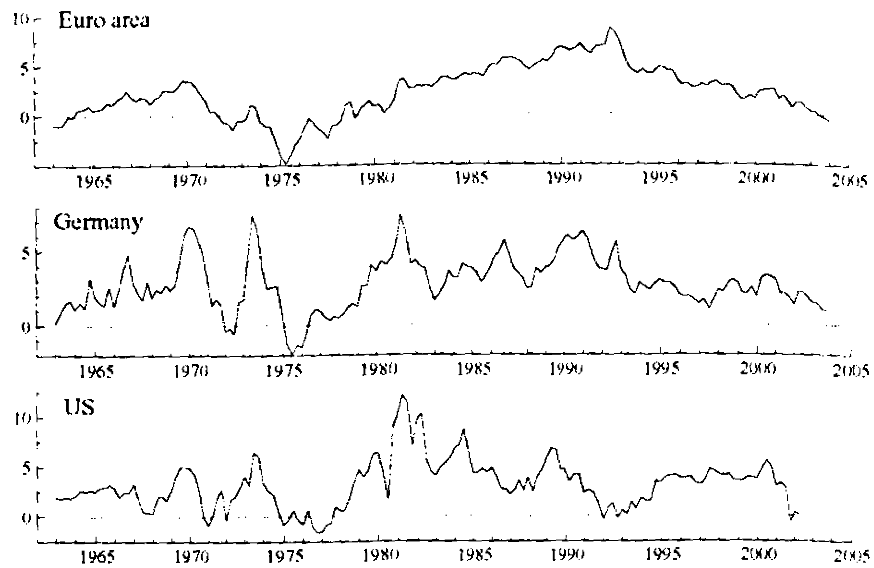

Figure 4.2: Real interest rate in three countries

4.3.2 Theoretical basis for the natural rate of interest

This paper takes a 'long-run' time-perspective and uses economic theory as a benchmark for determining the developements in the natural real interest rate. As recalled by LW, standard growth models imply that the natural real interest rate varies over time in response to shifts in preferences and output growth. r_t^* can be represented as the equilibrium interest rate derived from the Euler equation for consumption

$$r_t^* = \theta g_t + \rho_t \tag{4.1}$$

where r_t^* is the natural (or equilibrium) real interest rate, g_t is the (log of) per capita consumption growth, θ is the relative risk aversion[5], and ρ_t is the discount rate, or households time preference in their general utility function. Stokey and Rebelo (1995), Blanchard and Fischer (1989) and Barro and Sala-i-Martin (1990) have provided estimations of eq.(4.1). However, there is a high uncertainty surounding the measures of households risk aversion and time preference.

Additional determinants may as well contribute to the time-variation in the natural

[5]Equivalently, $1/\theta$ is the inter-temporal elasticity of substitution in consumption

rate.[6] For instance, a reduction in public saving might put upward pressure on the natural rate. Similarly, a change in the uncertainty about interest rates and future inflation might affect the saving ratio and in turn the natural real interest rate. To sum up, measuring the natural real interest rate directly on the basis of (4.1) is not straightforward, which argues in favour of a more complete structural model and an indirect estimation of r_t^*.

4.3.3 The Laubach & Williams model

The empirical framework suggested by LW is to run the Kalman filter on a system of equations to jointly estimate the natural real interest rate, potential output growth rate and the output gap. They propose a model of a neo-keynesian inspiration, that jointly characterises the behaviour of inflation and the output gap through modified IS and Phillips curves. As stated above, neo-keynesian models are not so much interested in the levels of variables composing these curves than in the deviations from equilibrium values. The main equations of the model are given by

$$a_y(L)\tilde{y}_t = a_r(L)\tilde{r}_t + \mathbf{a'x_t} + \varepsilon_{\mathbf{1,t}} \tag{4.2}$$

$$b_\pi(L)\pi_t = b_y(L)\tilde{y}_t + \mathbf{b'x_t} + \varepsilon_{\mathbf{2,t}} \tag{4.3}$$

where $\tilde{r}_t$ is the real interest rate gap, $\tilde{y}_t$ represent s the output gap defined as the difference between the (log) GDP y_t and (log) potential output y_t^* such that

$$\tilde{y}_t = 100(y_t - y_t^*) \tag{4.4}$$

π_t is consumer price inflation, and $\varepsilon_{1,t}$ and $\varepsilon_{2,t}$ are white noise errors, while a_y, a_r, b_π and b_y are polynomials in the lag operator L such that $a_y(L) = -\sum_{i=0}^{n} a_{y,i}L^i$, with $a_{y,0} = -1$. $\mathbf{x_t}$ is a vector composed of exogenous variables[7].

The laws of motion of the unobservable potential output and its trend growth rate are modelized by a local linear trend:

$$y_t^* = y_{t-1}^* + g_{t-1} + \varepsilon_{4,t} \tag{4.5}$$

$$g_t = g_{t-1} + \varepsilon_{5,t} \tag{4.6}$$

[6]See Bjorkstend and Karagedikli (2003), Cour-Thimann et al (2004) and Laubach and Williams (2003) for examples of factors that are not covered by the standard growth theory.

[7]Note that this vector has been dropped in the latest version of the program in order to simplify calculations and ease matrix manipulations. In earlier versions of the model, we had inserted a variable for relative energy prices and a dummy aiming at capturing the structural break induced by the 1990 German reunification. Both have been suppressed since the results were almost not affected.

where $\varepsilon_{4,t}$ and $\varepsilon_{5,t}$ are white noise errors.

Equation (4.1) is approximated with the following relationship for the natural real interest rate:

$$r_t^* = cg_t + z_t \tag{4.7}$$

where g_t is the unobservable trend growth rate of the economy, and c a parameter capturing the relative risk aversion. z_t represents other possible determinants of the natural rate of interest, such as households time preference, variations in public saving or uncertainty about interest rates.

In LW, z_t is either a stationary AR(2) process or a random walk. The measure of the random determinants of the natural real interest rate is obviously associated with a considerable amount of uncertainty. We face here a technical problem in that r_t^* is an unobserved variable, itself composed of two unobserved components. This difficulty has already been pointed out by Mésonnier & Renne (2004). In most of the specifications that we have tried, the results were highly sensitive to initial conditions and were often not reasonable. This is especially the case when z_t follows a random walk. Indeed, while the first element g_t is explicitly linked to the output through eq.(4.4) and (4.5), z_t is only defined through (4.8) below, which makes it more sensitive to small variations in the initial specifications of the model. To overcome this problem, we only consider the case where z_t follows a stationary AR process, and we restrict its variance σ_3^2 through a signal-to-noise ratio, λ_z –see the next section on this point. In addition, we claim that the possible elements composing z_t should be stationary, thereby making a random walk specification useless.

$$z_t = \alpha z_{t-1} + \varepsilon_{3,t} \tag{4.8}$$

(4.7), (4.8), (4.5) and (4.6) constitute the state (transitory) equations of our state-space model, while the IS curve (4.2) and the Philipps curve (4.3) constitute the observation equations (see Harvey (1989)). On this system, the Kalman filter is run twice. First in order to identify parameters by maximum likelihood. Second, in order to estimate the unobserved components r_t^*, y_t^*, g_t and z_t. The model can be written under its state-space form (see appendix):

$$\mathbf{y}_t = \mathbf{Z}\alpha_t + \mathbf{B}\mathbf{x}_t + \mathbf{G}\varepsilon_t \tag{4.9}$$

$$\alpha_{t+1} = \mathbf{T}\alpha_t + \mathbf{H}\varepsilon_t \tag{4.10}$$

4.3.4 Model estimation

The procedure follows different steps, in line with the recommendations of LW. The first one is to get a prior estimation of the output gap. For this purpose, we use a segmented linear trend with breaks in 1973 and 1993, as a proxy for potential output. This initial output gap is then used to estimate the coefficients of the simplified system by OLS: $a_y(L)\tilde{y}_t = a_r(r_{t-1} + r_{t-2}) + \varepsilon_{1,t}$ (for the IS equation) and $b_\pi(L)\pi_t = b_y(L)\tilde{y}_t + \varepsilon_{2,t}$ (for the Phillips curve). This allows to get adequate starting values for the maximum likelihood estimation of the coefficients.

In a second step, we consider a simplified system similar to the previous one, except that we estimate the coefficients by maximum likelihood and that we use the Kalman filter. The potential output y_t^* is treated as an unobserved component:

$$a_y(L)\tilde{y}_t = a_r(r_{t-1} + r_{t-2}) + \varepsilon_{1,t}$$

$$b_\pi(L)\pi_t = b_y(L)\tilde{y}_t + b_x x_t + \varepsilon_{2,t}$$

$$y_t^* = y_{t-1}^* + \bar{g} + \varepsilon_{4,t}$$

where $\tilde{y}_t$ is the output gap, r_t is the real interest rate, and $\bar{g}$ is a constant. Compare with (4.2)-(4.6).

The third step is dedicated to the median unbiased estimate of potential output growth variance, σ_5^2.[8] For this purpose, we use the estimate of y_t^* from the previous step in order to run the median unbiased technique of Stock & Watson (1998). The procedure works as follows: 1/ Regress for every date t the potential output growth[9] on a constant and a dummy with a break at time t. 2/ Compute the t-ratios corresponding to the coefficients of the dummies. 3/ Compute the Exponential Wald (EW) statistic[10]. It adds more or less the t-ratios obtained at every date. 4/ Compare the values obtained with that of Stock & Watson's table that maps these statistics to the value of median unbiased signal-to-noise ratios λ. Once we have found the adequate ratio λ_g,[11] it suffices to plug it into $\sigma_5 = \lambda_g \sigma_4$

[8] The reason for using this approach is that the bulk of the distribution of the parameters that control for the variance is often very close to zero. Consequently, the maximum likelihood estimates of these parameters are often statistically unsignificant, and are much below the median of the distribution. This would imply for instance that g_t would be constant.

[9] i.e. Δy_t^*, where y_t^* is computed from the second step system.

[10] such that $EW = \ln(\frac{1}{T}\sum_{t=1}^{T}\exp(s_t^2/2))$ where s_t is the t-ratio corresponding to a break at time t.

[11] We keep here the same notation as LW.

in order to get the adequate potential output. We provide below (appendix A.4.2) a sensitivity analysis of the model by taking different percentiles of the distribution of λ_g's, computed from 10.000 draws of the monte carlo simulation procedure used by Laubach & Williams.

As exposed above, the variance of z_t is set according to the signal-to-noise ratio $\lambda_z = \frac{a_r}{\sqrt{2}}\frac{\sigma_3}{\sigma_1}$.[12] We use once again the median unbiased estimator. Although this technique is only needed in theory when z is non-stationary, it should provide an adequate way to estimate λ_z, even with stationary processes. For this purpose, we compute on the euro area the monte carlo procedure of Laubach & Williams to get a distribution for λ_z. We use the median of this distribution as our baseline value. That is, we take $\lambda_z = 0.064$. The 5th percentile is equal to 0.046 and the 95th one to 0.076.

The final step estimates the whole system (4.2) to (4.8) by maximum likelihood[13], with the two ratios λ_g and λ_z imposed. We proceed in two steps: estimate the whole system first and store the IS curve output lag coefficients. Second, re-estimate the whole system with these coefficients fixed. For some reasons, the estimated output gap with this procedure is much more in line with the existing literature than in the case of a simple, one step estimation.

In order to identify the model, we have to restrict some parameters. For instance, the variance parameters were imposed to be strictly positive. This is a common practice in the literature. We also use some constraints that are specific to the model. For example, we impose $a_g = c.a_r \leq 0$, since there should be a mechanical negative relation between the growth rate g_t of potential output y_t^* and the output gap $\tilde{y}_t = y_t - y_t^*$. This in turn implies that the coefficient c is positive, which is also intuitive because this coefficient is supposed to capture consumers' relative risk aversion. Following LW, we also take a simple moving average of the first and second lags of $\tilde{r}_t$. This comes down to imposing the same coefficient on these two elements. We suppose in addition that the resulting coefficient a_r is less than or equal to zero, since the real rate gap should be countercyclical.

In practice, we follow Laubach and Williams (2003) and we assume that the polynomials $a_y(L)$ and $a_r(L)$ in equation (4.2) are of order 2, while $b_y(L)$ in (4.3) is simply of order

[12] $\sqrt{2}$ comes from the assumption that the output gap in equation (4.2) is determined by a moving average of the real interest gap of order 2. That is, $\tilde{y}$ is influenced by z_{t-1} and z_{t-2} through a single coefficient, a_r. See the following section.

[13] The BFGS procedure for numerical optimization is used for this purpose.

1. $b_\pi(L)$ is of order 3, but instead of taking the second and third lags of π_t, i.e. π_{t-2} and π_{t-3}, we take moving averages of the last three quarters of the first year and the whole previous year, such that : $\pi'_{t-2} = \sum_{i=2}^{4} \pi_{t-i}$ and $\pi'_{t-3} = \sum_{i=5}^{8} \pi_{t-i}$.

4.4 Results

4.4.1 Estimations of r_t^*

This section reports and discusses the estimation results. Table 1 shows parameter estimates for the euro area, Germany and the US. The estimates are generally similar to those reported by LW on US data.

As regards the Phillips curve, the null hypothesis that the coefficients of the inflation terms sum to one is not rejected by the data. The sum of the coefficients of the autoregressive components of the output gap lies between 0.83 and 0.93 in all countries. The effect of a change in the real interest rate gap on the output gap seems to be somewhat weaker in the euro area than in the US and Germany. The effect of a change in the output gap on inflation, on the other hand, seems to be slightly stronger in the euro area compared to the US and Germany.

Table 1 : Parameter estimates, baseline model

	Euro Area	*Germany*	*US*
	63q1-04q1	*63q1-04q1*	*61q1-02q2*
Variances			
λ_g	0.081	0.081	0
λ_z	0.064*	0.064*	0.064*
σ_{IS}	0.005	0.008	0.006
$\sigma_{Phillips}$	0.396	0.473	0.776
σ_{y*}	0.003	0.004	0.004
$\sigma_g = \lambda_g.\sigma_{y*}$	2.43×10^{-4}	3.21×10^{-4}	0
IS curve			
a_{y1}	0.70 (1.88)	0.47 (1.23)	1.63 (6.63)
a_{y2}	0.14 (1.81)	0.36 (1.23)	-0.70 (6.75)
a_r	-0.056 (2.42)	-0.172 (1.47)	-0.18 (1.96)
c	0.880	0.653	1.179
Phillips curve			
$b_{\pi1}$	1.18 (6.25)	1.07 (4.65)	0.77 (3.04)
$b_{\pi2}$	-0.28 (5.34)	-0.14 (4.00)	0.13 (2.60)
$b_{\pi3} = 1 - (b_{\pi1} + b_{\pi2})$	0.1*	0.07*	0.09*
b_y	0.051 (9.31)	0.041 (7.73)	0.103 (18.01)

t statistics in parentheses

*: imposed coefficient

Regarding the estimate of the natural real interest rate in the euro area, Figure 4.3 reveals a very high uncertainty around the estimate of the *level* of r^*. However, as indicated in the sensitivity analysis of the appendices, we argue that this is essentially due to the uncertainty surrounding the coefficient c. If we impose a constraint on the possible range of values this coefficient can take, the range of estimates of the level of the natural real interest rate diminishes. For almost all signal-to-noise ratios in the model (having fixed c to its baseline estimate), the estimated natural real interest rate seems to have been higher in the 1960s and early 1970s than in the 1990s and 2000s. Another interesting result is that the natural real interest rate seems to have been higher in 1990, when the reunification

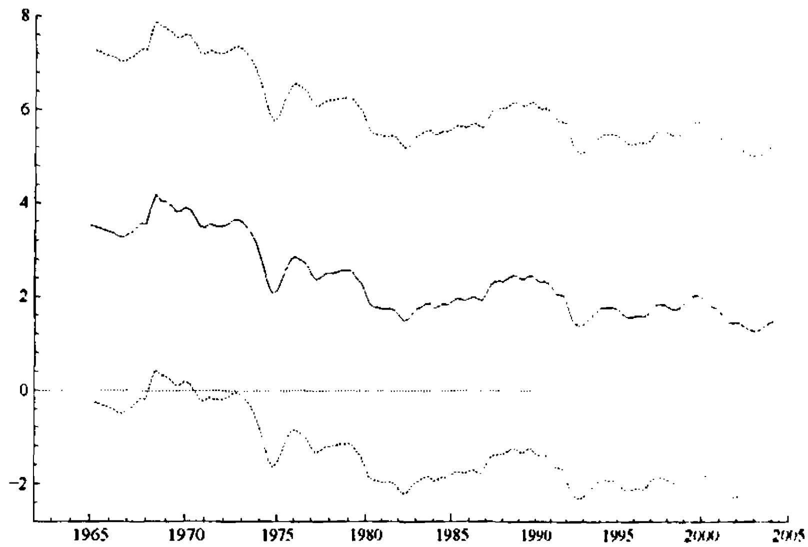

Figure 4.3: Natural real interest rate r_t^* estimate and ±2 s.d.

of East and West Germany took place, compared to the period after the Stage Three of the EMU. Furthermore, the estimated natural real interest rate was also lower in 2004 compared to the start of Stage Three of the EMU in 1999. Finally, our baseline estimate suggests that the natural real interest rate has declined from around 4% in the 1960s to less than 2% in 2004 (figure 4.3). In the model, this decline in the natural real interest rate in the euro area owes in particular to a fall in the estimated trend growth rate of the economy (see Figure 4.4 below)[14].

Figure 4.5 shows our baseline estimate of the output gap. The estimate is consistent with the common held view that monetary policy was loose in the 1970s, contributing to a positive output gap and a persistent high level of inflation for most of the decade. Moreover, in the 1980s and early 1990s, as monetary policy authorities in many countries pursued a tight monetary policy oriented towards disinflation[15], the output gap turned

[14] Cour-Thimann et al (2004) argue for the view that increases in government debt in the 1980s and higher exchange rate risk premia in the early 1990s might have put an upward pressure on the natural real rate in the euro area. These arguments would imply that our estimate of the natural real interest rate in this period is somewhat low.

[15] See for instance Taylor (1992) for a description of the disinflation policy in the US.

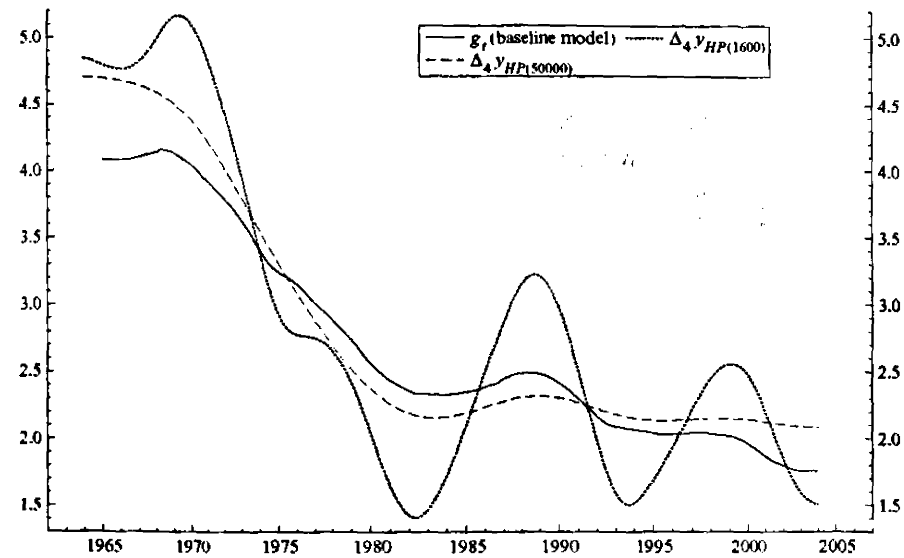

Figure 4.4: Estimated potential output growth g_t and fourth differences of HP-filtered output

negative in the 1980s. Except for a small positive output gap in the beginning of the 1990s, the output gap remained negative until the start of Stage Three of EMU. Interestingly, Figure 4.6 shows that inflation began to decline almost immediately after the estimated real interest rate gap turned positive in the 1980s.

Starting with the 1960s, at a time where financial markets in some euro area countries were still heavily controlled, the low real interest rate did not seem to have put upward pressure on inflation as one would expect. An interpretation could be that any measure of the natural rate in the 1960s is particularly uncertain, or that a relatively wide range of possible real rate levels would be in line with maintaining price stability in this decade.

For the 1970s, there are reasons to believe that the estimated real interest rate gap, which is negative on average, draws a quite reasonable picture of the monetary policy stance. While it is clear that the oil price increases contributed to the rise in inflation in the 1970s, they were arguably followed by inadequate monetary policy (see the box entitled "Current euro area interest rates from a historical perspective" on page 25 of the September 2003 issue of the ECB Monthly Bulletin and the box entitled "Lessons to be drawn from the oil price shocks of the 1970s and early 1980s" on page 21 of the November 2000 issue of the Monthly Bulletin). Nominal interest rates rose, on average, to very high levels in a historical perspective in the 1970s. However, monetary policy in many euro area countries did not initially react to the necessary extent and was too expansionary in the whole decade. The loose monetary policy stance in the 1970s contributed therefore to a positive output gap for most of the decade, causing a persistent high level of inflation.

Regarding the 1980s and early 1990s, there are reasons to believe, in our opinion, that positive real rate gaps drawn by our estimates in this period are indeed reasonable. Giammarioli and Valla (2003) estimates of the real interest rate gap for the 1980s and 1990s are comparable to the results in this paper. A common view is that monetary policy authorities in many euro area countries, but also in other countries like the US and UK, pursued a tight monetary policy oriented towards disinflation –see Taylor (1992). Comparing chart 4.5 and 4.6, our estimates suggest that this disinflation policy contributed to the negative output gap in the 1980s. Except for a small positive outout gap in the beginning of the 1990s, it remained negative until the start of Stage Three of EMU. A noticeable feature is that the measure of output gap differs from that obtained from ad-hoc filtering. For instance, it was on average lower from the 1980s. This could interestingly be related to disinflationary policies that started roughly in the same period. The estimated output gap could be seen as the one taking into account the influence of monetary policy on the real economy. Such an output gap might have an important policy impact if it was taken into account.

As regards inflation, it started to fall just after the real interest rate become positive in the 1980s and, after a small increase at the end of the decade, it continued to decline in the 1990s (figure 4.6).

Figure 4.7 compares the baseline estimates of the natural real interest rates for the euro area, Germany and the US. Evidently, while the natural real interest rate in the euro area has declined over the sample, the natural real interest rate in the US has been more stable around its long term average. The estimates also indicate that the natural real interest rate is lower in the euro area, and in particular in Germany, than in the US.

It is important to stress that all estimates of the natural real interest rate are very imprecise and that caveats are associated with all estimation methods. Regarding the pitfalls with the approach taken in this paper, the estimation results are very sensitive to

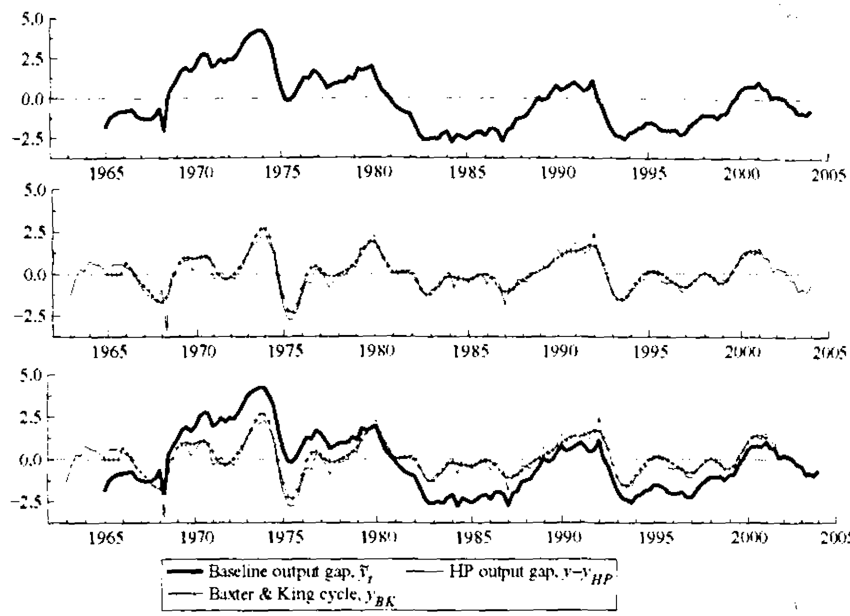

Figure 4.5: Comparing output gaps: Baseline estimate $\tilde{y}_t$ and two standard filtering methods

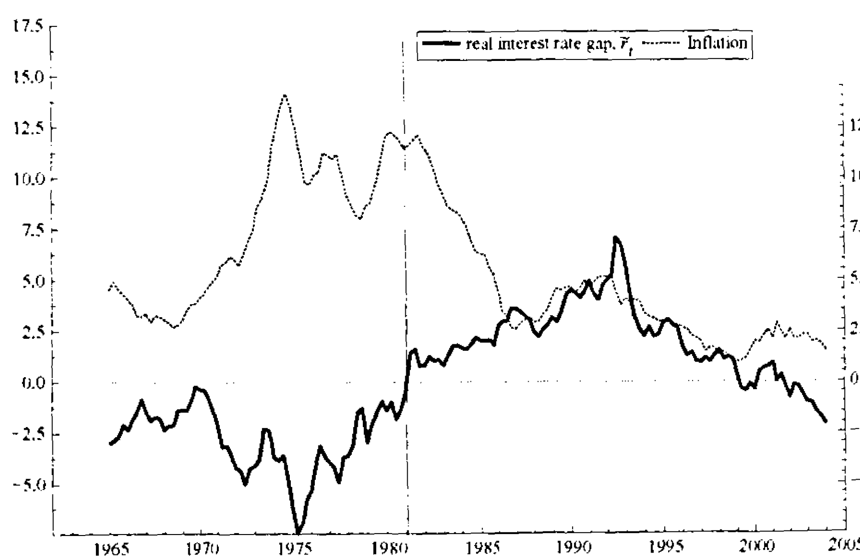

Figure 4.6: Natural real interest rate gap $\tilde{r}_t$ ($= r_t - r_t^*$) and inflation rate, euro area.

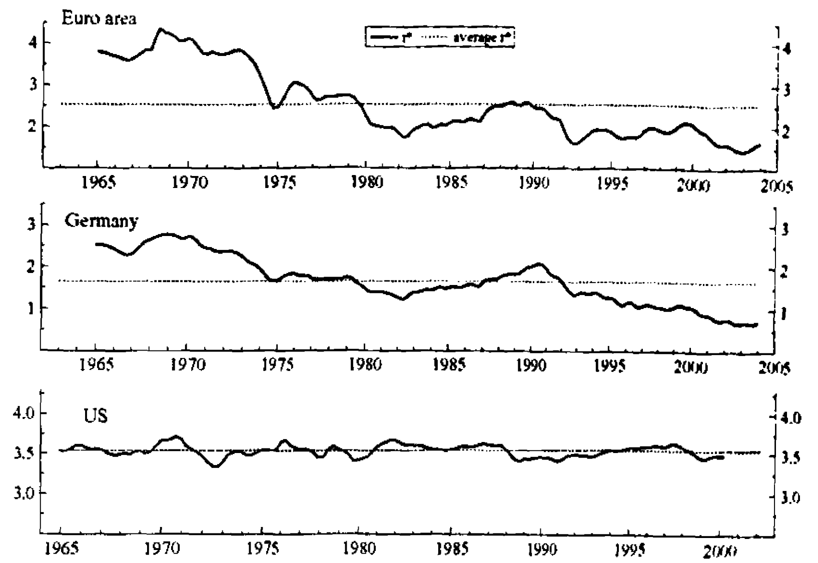

Figure 4.7: Estimations of r_t^* for the euro area, Germany and the US

initial specifications of the model and the selection of starting values for the parameters. A second aspect concerns the measurement of time-variations in preferences (see equation 4.1). Third, non-textbook factors that may contribute to the time variation in the natural rate are treated arbitrarily. Within the empirical framework of this paper, variable z_t is supposed to represent all other factors than trend output growth to explaining the developments in the natural real interest rate. Arguably, the preciseness of this measure is doubtful, which could make the estimates difficult to interpret.

4.4.2 Some properties of the real interest gap in the euro area

We now examine some statistical properties of the euro area model, focusing on simple statistics that describe the relationship between the real interest rate gap and inflation. Table 2 and 3 report standard deviations and correlation of selected variables used in this analysis, namely log output y_t, log potential output y_t^*, the output gap $\tilde{y}_t$, the actual and the natural real rate and the real rate gap (r_t, r_t^* and $\tilde{r}_t$) and inflation π_t. A notable feature of the reported statistics is that the correlation between the actual real interest rate and the real interest rate gap are high and their standard deviations are roughly identical.

In other words, the variation in the real interest rate is not primarily related to variation in the natural real interest rate. This is consistent with the results in Giammarioli and Valla (2003) for the euro area, Neiss and Nelson (2003) for the UK and Laubach and Williams (2003) for the US, but stands against the results of Cour-Thimann (2004) for the euro area.

Table 2:
Standard deviations, euro area

y_t	0.35
y_t^*	0.30
$\tilde{y}_t$	1.72
π_t	3.50
r_t	2.55
r_t^*	0.80
$\tilde{r}_t$	2.93

Table 3 : Correlation coefficients, euro area

	$k=0$	$k=1$	$k=2$	$k=3$	$k=4$
$Corr(r_t, \tilde{r}_{t-k})$	0.98				
$Corr(\pi_t, \tilde{r}_{t-k})$	-0.47	-0.48	-0.50	-0.53	-0.55
$Corr(\tilde{y}_{tt}, \tilde{r}_{t-k})$	-055	-0.57	-0.59	-0.60	-0.62

Interestingly, the correlation between inflation and the real rate gap is strongly negative at all lags. This indicates that the developments in euro area inflation since 1960 is, in part, related to the evolution of the real interest rate gap[16]. In line with this hypothesis we find that the real rate gap is also strongly and negatively correlated with the output gap. Next, we investigate the leading indicator properties of the real interest rate gap on inflation. Following Neiss and Nelson (2003), we estimate:

$$\pi_t = a + b_1\pi_{t-1} + b_2(r_{t-k} - r^*_{t-k}) + \varepsilon_t$$

[16] Neiss and Nelson (2003) and Giammaioli and Valla (2003) report similar correlation coefficients and draws similar conclusions for the UK and euro area respectively.

where annual inflation π_t is regressed on past inflation and lagged values of the real interest rate gap. The regression on the 1962:1 - 2003:1 sample is summarised in table 4 (numbers in parentheses are standard errors). The results indicate that lagging the real interest rate gap 3 to 5 quarters yields statistically significant parameter estimates when added to an autoregression for inflation. The long lags seem consistent with the common view that monetary policy affects inflation after a significant delay. This simple exercise suggests that the estimated real interest rate gap may contain valuable information about future inflation.

Table 4 : Parameter estimates, euro area

	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$
a	0.08	0.10	0.13	0.15	0.18
b_1	0.98	0.98	0.97	0.97	0.97
b_2	-0.02 (-1.53)	-0.03 (-1.97)	-0.04 (-2.85)	-0.05 (-3.07)	-0.06 (-3.70)
R_2	0.98	0.98	0.98	0.98	0.98

4.5 Conclusion

This paper estimates the natural real interest rate for the euro area, considering the currency union as a single entity over the period 1963 to 2004, and for Germany. Following closely the methodology suggested by Laubach and Williams (2003), we apply the Kalman filter on a small scale macroeconomic model, which encompasses a Phillips curve and an IS curve. This allows us to estimate the natural real interest rate, potential output and trend growth rate of the two economies simultaneously. Overall, the results are quite comparable with the original results of Laubach and Williams (2003) using US data.

According to our baseline results, the fluctuations in the natural real interest rate have been relatively low since 1963. The natural rate has declined gradually over the past 40 years, from an estimate of around 4% in the 1960s to slightly less than 2% in 2004. The real interest rate gap is relatively persistent over longer periods, with low short-term fluctuations.

Regarding the output gap, the length of the business cycle's booms and busts are in line with the consensus view in the business cycle literature. However, this model produces an output gap that is influenced by the real interest rate gap. Its average level is negative in periods of restrictive monetary policy and positive in periods of monetary laxism. This is interesting since it might be taken as an indicator of the degree at which the central bank policy influences the real economy. In the 1970s, when inflation became high, the real interest rate gap was negative and the output gap was positive, on average. In the 1980s and the 1990s, when inflation fell to lower levels, the real interest rate gap was positive and the output gap was negative, on average.

Simple empirical tests also suggest that the estimated real interest rate gap is negatively correlated with the output gap and inflation. Furthermore, the tests show that it may contain valuable information about future inflation in the euro area. The general caveats associated with interpreting estimates of the natural rate, which are highly uncertain, also applies for this paper.

Appendix

A.1 Appendix for Chapter 1

A.1.1 Description of the dating procedures

- **Bry & Boschan (1971)**

1. Determination of extremes and substitution of values.
2. Determination of cycles in 12-month moving average (extremes replaced).
   (a) Identification of points higher (or lower) than 5 months on either side.
   (b) Enforcement of alternation of turns by selecting highest of multiple peaks (or lowest of multiple troughs).
3. Determination of corresponding turns in the Spencer curve (extremes replaced).
   (a) Identification of highest (or lowest) value within $\pm$ 5 months of selected turns in 12-months moving average.
   (b) Enforcement of minimum cycles of duration of 15 months by eliminating lower peaks and higher troughs of shorter cycles.
4. Determination of corresponding turns in short-term moving average of 3 to 6 months, depending on MCD (months of cyclical dominance).
   (a) Identification of highest (or lowest) value within $\pm$ 5 months of selected turns in Spencer curve.
5. Determination of turning points in unsmoothed series.
   (a) Identification of highest (or lowest) value within $\pm$ 4 months, or MCD term, whichever is larger, of selected turn in short-term moving average.
   (b) Elimination of turns within 6 months of beginning and end of series.
   (c) Elimination of peaks (or troughs) at both ends of series which are lower (or higher) than values closer to end.
   (d) Elimination of cycles whose duration is less than 15 months.
   (e) Elimination of phases whose durations is less than 5 months.

(f) Statement of final turning points.

- **Artis et al. (1997)**

1. Determination of extreme values (those for which the log-change with respect to adjacent month is greater than 3.5 standard errors of the log-differenced series).
2. Determination of cycles in the series smoothed with an MA(7).
   (a) Identification of peaks/troughs within ± 12 months.
   (b) Enforcement of alternation of turning points (same as before).
3. Determination of turning points on unsmoothed series.
   (a) Points higher/lower within ± 12 months.
   (b) Enforcement of alternation of peaks and troughs.
   (c) Identification of flat segment (those for which it is not possible to say if the phase is "expansionary" or "contractionary").
   (d) Identification and exclusion of outliers from the first set of turning points.
   (e) New enforcement of alternation.
   (f) Identification of short cycles (less than 15 months).
   (g) Enforcement of an amplitude of the phases superior to one standard error of the (log) changes.
4. Comparison of the turning points taken from the smoothed and the unsmoothed series and elimination of the points that do not correspond to similar turns (± 5 months of the moving average)
5. statement of the final set of turning points.

- **Procedure used here**

1. Elimination of outliers: same as AKO.
2. Determination of cycles in MA(7).
   - Identification of peaks/troughs within ± 12 months.

3. Determination of turning points on unsmoothed series.

   (a) Points higher/lower within ± 12 months.

   (b) Identification of short cycles (less than 15 months).

   (c) Enforcement of an amplitude of the phases superior to one standard error of the (log) changes.

4. Comparison of turning points: same as AKO.

   (a) Enforcement of alternation.

   (b) Statement of the final set of turning points.

A.1.2 Turning points dates

Table B.1 : Turning points dates

ECRI	Own	BBW	AKO	ECRI	Own	BBW	AKO	ECRI	Own	BBW	AKO
Austria				*Belgium*				*Finland*			
Peaks				*Peaks*				*Peaks*			
Aug-74	**Jun-74**	**Jun-74**		n.a		Dec-70		n.a	Jul-74	Jul-74	
Feb-80	**Dec-79**	**Dec-79**			Jan-74	Jan-74	Apr-74			Jan-82	
		Dec-82			Feb-77	Feb-77	Oct-76		Jan-90	Jan-90	
		Mar-86				Dec-79	Dec-79		Dec-00		
	Dec-90	Dec-90				Jul-86					
Apr-92											
May-95		**Jul-95**			Nov-90	Nov-90	Mar-90				
Jan-01	**Nov-00**	**Feb-01**			Feb-92	Feb-92					
						Feb-95					
						Jul-98					
					Nov-00	Nov-00					
Troughs				*Troughs*				*Troughs*			
Jun-75	Oct-75	Oct-75		n.a		May-71		n.a	Sep-75	Sep-75	
		Jul-81			Aug-75	Aug-75	Jul-75			Jul-82	
Jan-83	**Dec-82**	**Dec-82**					Sep-77		Oct-91	Oct-91	
		Jan-87			Jan-79	Jan-79			Apr-01	Dec-01	
Jun-93	Dec-92	**Jun-93**					Dec-80				
Mar-96											
Dec-01	**Mar-02**	**Mar-02**				Apr-84					
					Jan-87	Jan-87					
					Feb-91		Aug-91				
					Nov-93	Nov-93					
						Feb-96					
						Feb-99					
					Oct-01	Oct-01					
France				*Germany*				*Greece*			
Peaks				*Peaks*				*Peaks*			
		Apr-64		Mar-66	**Mar-66**	**Mar-66**	**Mar-66**	n.a	Feb-74	Feb-74	
Jul-74	**Jul-74**	**Aug-74**	**Aug-74**	Aug-73	**Aug-73**	**Aug-73**	**Aug-73**		Apr-80	Apr-80	
	Sep-76	Jan-77	Jan-77	Jan-80	**Dec-79**	**Dec-79**	**Dec-79**			May-82	
Aug-79	**Jul-79**	**Aug-79**	**Aug-79**			Jul-86				Dec-85	
Apr-82	Dec-81	Dec-81	Dec-81	Jan-91	**Jan-91**	**Jan-91**	**Jun-91**		Feb-90	Feb-90	
		Apr-86			Dec-94	Dec-94			Dec-00	Dec-00	
	Jan-91	Jan-91				Jul-98				Apr-02	
Feb-92			**Apr-92**	Jan-01	**Feb-01**	**Feb-01**					
	Jun-95	Jun-95				Nov-02					
	Jan-01	Jan-01									
Troughs				*Troughs*				*Troughs*			
		Jan-65		May-67	**May-67**	**May-67**	**May-67**	n.a	Jul-74	Jul-74	
Jun-75	**May-75**	**May-75**	**May-75**	Jul-75	**Jul-75**	**Jul-75**	**Jul-75**			Apr-81	
	Dec-77	Dec-77	Dec-77	Oct-82	**Nov-82**	**Nov-82**	**Nov-82**		May-83	May-83	
Jun-80	Nov-80	Apr-81	Nov-80			Jan-87				Jul-87	
	Aug-82	Aug-82	Aug-82		Jul-93	Jul-93			Jul-93	Jan-93	
Dec-84				Apr-94					Dec-01		
Aug-93	**Aug-93**	**Aug-93**			Oct-95	Oct-95				Feb-03	
	Dec-95	Dec-95			Nov-01	Nov-01					
	May-03	May-03		Aug-03							

source : OECD
na : not available.
Bold : no more than 3 months of difference with ECRI dates.
nb: the dataset of AKO stops in December 1993.

Table A.1:

Table B.1 (II)

ECRI	Own	BBW	AKO	ECRI	Own	BBW	AKO	ECRI	Own	BBW	AKO
Italy				*Luxembourg*				*Netherlands*			
Peaks				*Peaks*				*Peaks*			
Jan-64	**Jan-64**	**Jan-64**	**Jan-64**	n.a	Feb-65	Feb-65	Feb-65	n.a	Aug-74	Aug-74	Aug-74
		Jul-69			Jan-70	Mar-70	Mar-70			Sep-76	Sep-76
Oct-70		Jan-71			Aug-74	Aug-74	Aug-74		Nov-79	Nov-79	Mar-80
Apr-74	**Jun-74**	**Jun-74**	**Jun-74**		May-76	May-76	May-76			Jan-85	
	Jan-77	Jan-77	Jan-77		Dec-79	Dec-79	Dec-79		Jan-87	Jan-87	Jan-87
May-80	**Mar-80**	**Mar-80**	**Mar-80**			Feb-82			Feb-91		Feb-91
	Dec-89	Dec-89	Dec-89		Dec-84	Dec-84	Oct-85			Jan-92	
Feb-92		**Feb-92**			Jun-90	Jun-90				Dec-95	
	Dec-95	Dec-95					May-92		Apr-01		
	Oct-97	Oct-97			Aug-95	Aug-95				Jun-02	
	Dec-00	Dec-00				Feb-98				Feb-04	
		Jul-02			Dec-00						
Troughs				*Troughs*				*Troughs*			
Mar-65	Aug-64	Aug-64	Aug-64	n.a	Aug-67	Aug-67	Aug-67	n.a	Aug-75	Aug-75	Aug-75
Aug-71					Oct-70	Oct-70	Oct-70			Nov-77	May-78
Apr-75	**Apr-75**	**Apr-75**	**Apr-75**		Aug-75	Aug-75	Aug-75		Nov-82	Nov-82	Nov-82
	Nov-77	Nov-77	Jun-77		Dec-76	Dec-76	Dec-76			Mar-86	
May-83	**May-83**	**May-83**	**Jun-83**		Apr-81	Apr-81	Apr-81		Apr-88	Apr-88	Apr-88
	Apr-91	Apr-91				Dec-82			Dec-92	Jun-93	
Oct-93		**Jul-93**			Feb-85		Oct-85		May-03	May-03	
	Dec-96	Dec-96			Aug-93	Aug-93					
						May-96					
	Dec-98	Dec-98			Jul-98	Jul-98					
		Nov-01									
	May-03	May-03									
Portugal				*Spain*				*Denmark*			
Peaks				*Peaks*				*Peaks*			
n.a		Apr-66			Aug-74	Aug-74	Aug-74	n.a	Apr-76	Aug-76	
	Jan-74	Mar-74		Mar-80		Aug-79			Jul-86	Jul-86	
	Nov-84	Nov-84				Jul-89	Jan-90			Nov-90	
	Aug-90	Aug-90		Nov-91		**Dec-91**	**Oct-91**		Jan-92		
		Nov-99			May-95	May-95			Jan-95	Jan-95	
	Apr-02	Nov-01			Nov-00	Nov-00			May-02		
		Sep-03								Apr-04	
Troughs				*Troughs*				*Troughs*			
n.a		Feb-67			Apr-75	Apr-75	Aug-75	n.a		Dec-74	
	Aug-75	Aug-75				Aug-82			Apr-77	Apr-77	
	Sep-85	Sep-85		May-84					Dec-87	Dec-87	
	Oct-93	Oct-93				Mar-91	Mar-91		May-93	May-93	
		Apr-00		Dec-93		Apr-93			Jan-96	Jan-96	
		Mar-03			Apr-96	Apr-96					
					Dec-01	Dec-01					

Table B.1 (III)

Sweden

ECRI	Own	BBW	AKO
Peaks			
Oct-70	Jul-70	Feb-71	
	Jul-74	Jul-74	
Jun-75			
Feb-80		**Mar-80**	
	Aug-85	Aug-85	
Jun-90		**Jun-90**	
		Nov-00	
		May-04	
Troughs			
Nov-71	**Aug-71**	**Aug-71**	
Nov-77	Jul-78	Jul-78	
		Aug-82	
Jun-83			
Jul-93	Dec-92	Dec-92	
		Apr-01	

UK

ECRI	Own	BBW	AKO
Peaks			
	Mar-66	Mar-66	Jul-66
	Oct-70	Jan-71	Jan-71
Sep-74	**Jun-74**	**Jun-74**	**Jun-74**
Jun-79	**Jun-79**	**Jun-79**	**Jun-79**
	Jan-84	Jan-84	Jan-84
May-90	**Jun-90**	**Jun-90**	**Jun-90**
		Jun-98	
		Dec-00	
		Jun-04	
Troughs			
	Nov-66	Nov-66	Nov-66
	Feb-72	Feb-72	Feb-72
Aug-75	**Aug-75**	**Aug-75**	**Aug-75**
May-81	**May-81**	**May-81**	**May-81**
	Aug-84	Aug-84	Aug-84
Mar-92		**Aug-91**	
		Feb-99	
	Jun-02		
		May-03	

Norway

ECRI	Own	BBW	AKO
Peaks			
n a		May-68	
	Oct-74	Nov-74	
	Jan-77	Jan-77	
	May-79	Oct-79	
	Aug-87	Aug-87	
	Nov-89	Aug-90	
		May-98	
		Oct-00	
	Apr-02	Apr-02	
		Jun-04	
Troughs			
n a		Feb-69	
	Dec-75	Dec-75	
	May-78	May-78	
	Mar-83	Mar-83	
	Aug-88	Aug-88	
	Dec-91	Dec-91	
		May-99	
		May-01	
	Jun-03	Jun-03	

Switzerland

ECRI	Own	BBW	AKO
Peaks			
		Feb-68	
Apr-74			
	Feb-75	Feb-75	
		Apr-77	
Sep-81			
	Apr-85	Jun-85	
Mar-90			
	Mar-93	Jul-93	
Dec-94	**Nov-94**	**Mar-95**	
Mar-01	Oct-00	**Jan-01**	
Troughs			
	Aug-68	Aug-68	
Mar-76		Nov-76	
	Sep-78	Sep-78	
Nov-82			
	Apr-87	Apr-87	
Sep-93	May-94	May-94	
Sep-96			
	May-99	May-99	
Mar-03	Nov-02		

Japan

ECRI	Own	BBW	AKO
Peaks			
Nov-73	**Jan-74**	**Jan-74**	**Jan-74**
		Feb-80	
	Oct-81	Nov-81	Nov-81
	May-85	May-85	May-85
	May-91	May-91	May-91
Apr-92			
Mar-97	**May-97**	**May-97**	
Aug-00	Dec-00	Dec-00	
Troughs			
		Dec-62	
Feb-75	**Mar-75**	**Mar-75**	**Mar-75**
		Aug-80	
	Oct-82	Oct-82	Oct-82
	Aug-86	Aug-86	Aug-86
Feb-94	**Jan-94**	**Jan-94**	
Jul-99	**Apr-99**	Dec-98	
	Nov-01	Nov-01	
Apr-03			

US

NBER	Own	BBW	AKO	ECRI
Peaks				
Dec-69	Aug-69	**Oct-69**	**Oct-69**	**Dec-69**
Nov-73	**Oct-73**	**Nov-73**	**Nov-73**	**Nov-73**
Jan-80	May-79	Jun-79	**Mar-80**	**Jan-80**
Jul-81	**Jul-81**	**Jul-81**	**Jul-81**	**Jul-81**
	Apr-89	Apr-89	Apr-89	
Jul-90		**Sep-90**		**Jul-90**
	May-00	Jun-00		
Mar-01				**Mar-01**
	Jun-02	Jun-02		
Troughs				
Nov-70	**Nov-70**	**Nov-70**	**Nov-70**	**Nov-70**
Mar-75	**Mar-75**	**Mar-75**	**Mar-75**	**Mar-75**
Jul-80	**Jul-80**	**Jul-80**	**Jul-80**	**Jul-80**
Nov-82	**Dec-82**	**Dec-82**	**Dec-82**	**Nov-82**
		Oct-89		
Mar-91	**Mar-91**	**Mar-91**	**Mar-91**	**Mar-91**
Nov-01	**Dec-01**	**Dec-01**		**Nov-01**
	Apr-03	Apr-03		

Bold : no more than 3 months of difference with NBER dates.

A.1.3 Typical cycles

The average growth rate of each sub-period can be represented as follows:

$$g^{p}_{k,s} = \frac{1}{N(t_{k,(s+1)_i} - 1 - t_{k,s_i})} \sum_{i=1}^{N} \sum_{t=t_{k,s_i}}^{t_{k,(s+1)_i}-1} \frac{y_{t_{k,s_i}} - y_{t_{k,s_i}-1}}{y_{t_{k,s_i}-1}}, \; p = \text{expansion, recession.}$$

Where N is the number of phases of the p-type, $y_{k,t}$ is the index of country k at time t, t_{k,s_i} is the date of beginning of the subperiod s of the phase i for country k.

Similarly, the average duration is

$$d^{p}_{k,s} = \frac{1}{N} \sum_{i=1}^{N} \left(t_{k,(s+1)_i} - 1 - t_{k,s_i}\right)$$

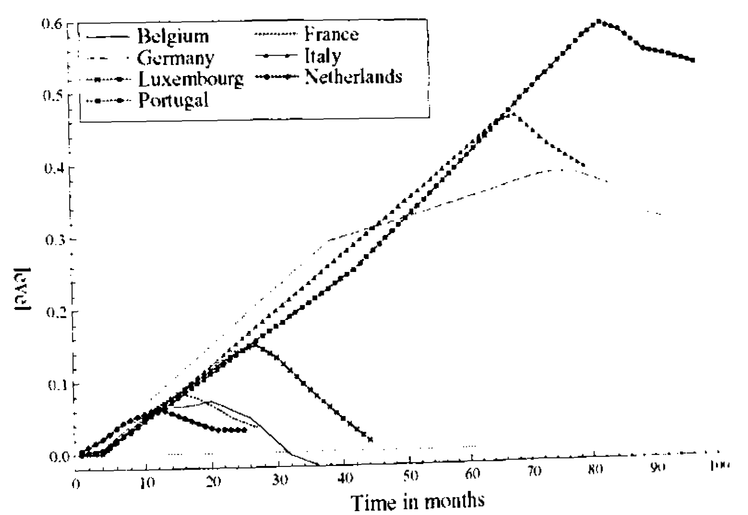

Figure A.8: Euro group - first period

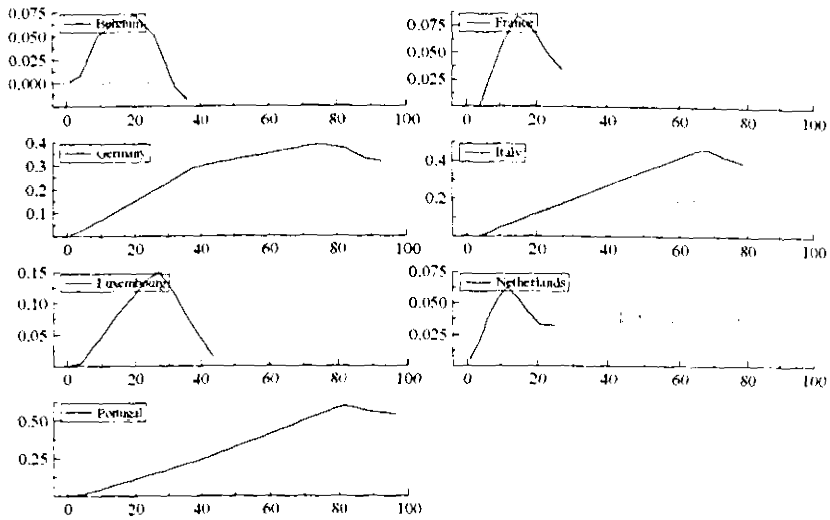

Figure A.9: Euro group (details) - first period

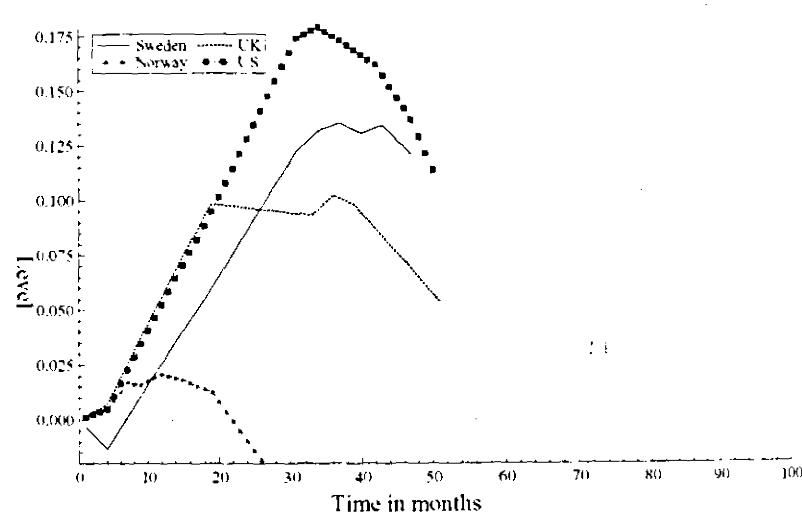

Figure A.10: 'EU without Euro' group - first period

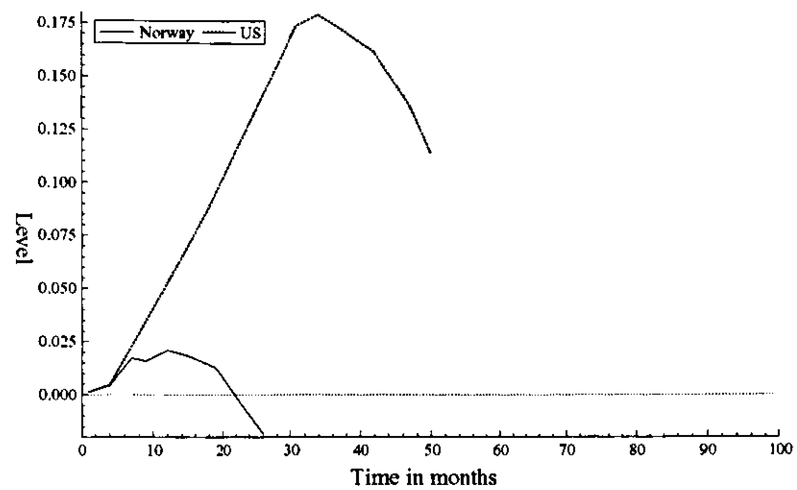

Figure A.11: non-EU group - first period

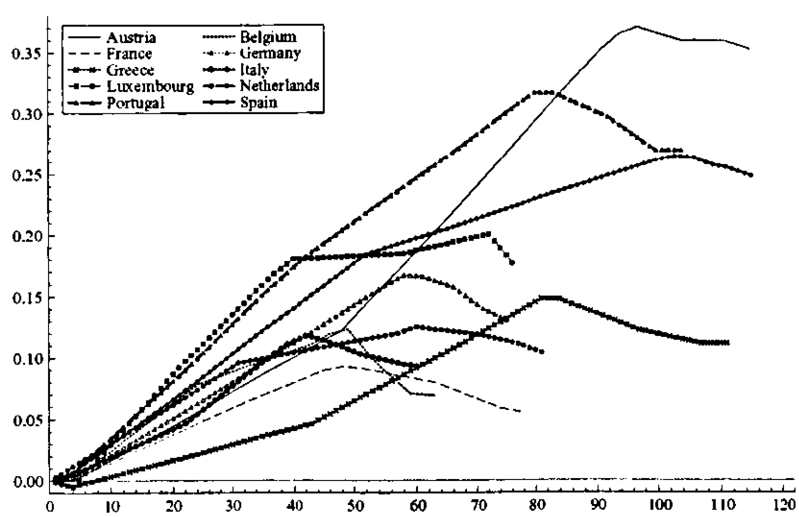

Figure A.12: Euro group - second period

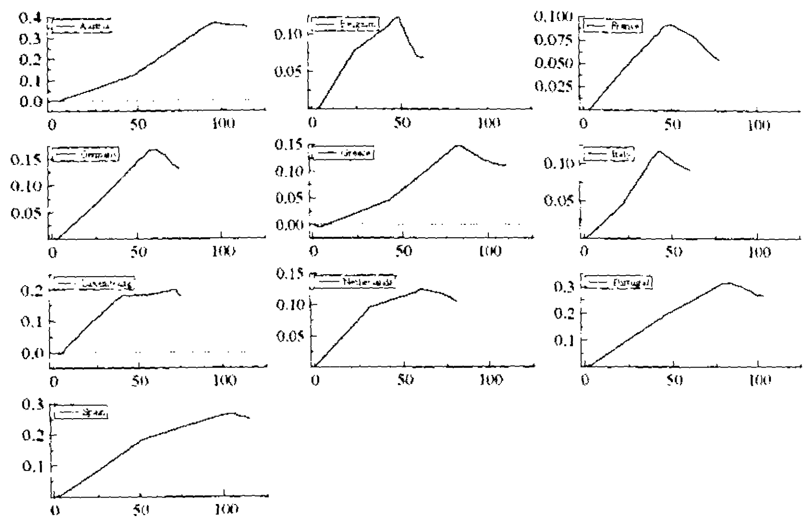

Figure A.13: Euro group (details) - second period

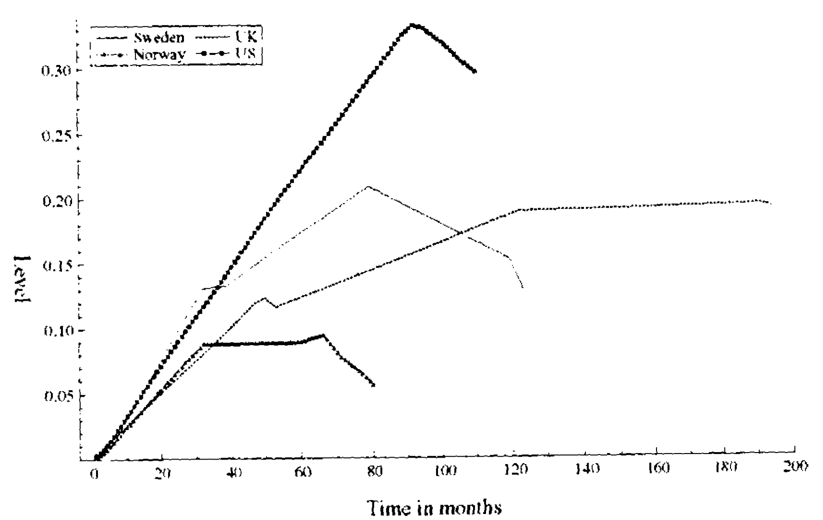

Figure A.14: 'EU without Euro' group - second period

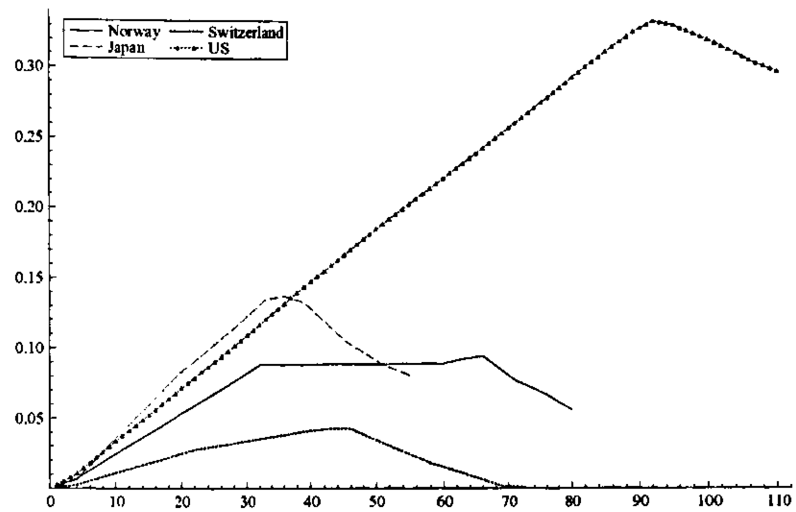

Figure A.15: non-EU group - second period

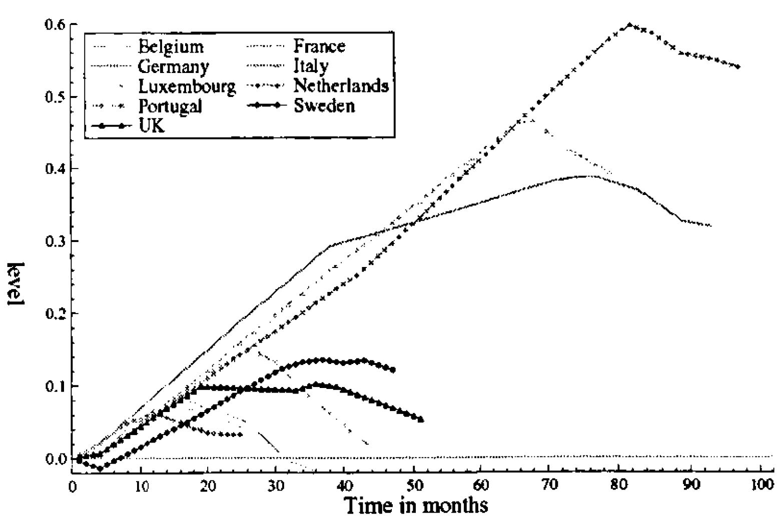

Figure A.16: Whole EU group - First period

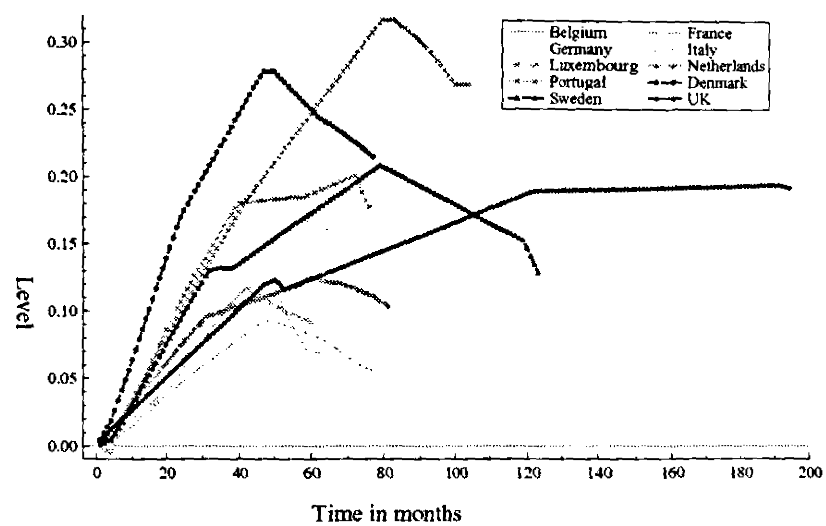

Figure A.17: Whole EU group - Second period

A.1.4 t-ratio tests

First period

Harding and Pagan's test using DenHaan and Levin robust t-ratio

	Euro group											EU without Euro			Externals			
	Au	Be	Fi	Fr	Ge	Gr	It	Lux	Net	Po	Sp	De	Swe	UK	No	Swi	Ja	US
Au		-	***	-	-	-	-	-	**	**	**		-	*	*	-	*	-
Be	-		-	-	-	-	-	-	-	*	-	-	-	-	*	-	-	-
Fi	***	-		*	*	-	-	-	***	**	**		-	*	*	-	*	-
Fr	*	*	*		-	-	**	-	*	-	*	-	**	-	**	-	-	-
Ge	-	-	-	-		-	-	-	-	**	-		-	-	-	-	*	-
Gr	-	-	-	-	**		-	-	-	***	-		-	-	-	-	***	**
It	*	*	*	***	-	-		-	*	*	**	-	*	-	**	-	*	-
Lux	-	-	-	-	*	-	-		-	-	-	*	-	-	-	-	-	-
Net	***	-	***	*	*	-	-	-		***	**		-	*	*	-	*	-
Po	**	-	**	-	*	-	-	-	**		*		-	-	-	-	**	-
Sp	***	-	***	**	*	-	**	-	***	**			-	*	-	-	**	*
De		-		-			-	-					-		-	-		
Swe	-	-	-	-	-	-	-	-	-	-	-	-		-	*	*	-	-
UK	**	-	**	-	*	-	-	-	**	*	-		-		-	-	-	-
No	*	*	*	**	-	-	*	-	*	-	-	-	**	-		*	-	-
Swi	-	-	-	-	-	-	-	-	-	-	-	-	-	-	-		-	-
Ja	*	-	*	-	*	**	-	-	*	***	**		-	-	-	-		**
US	-	-	-	-	-	-	-	-	-	*	-		-	-	-	-	**	

Significance levels:
*** : 1%
** : 5%
* : 10%

- : non rejection
Blank: test could not be conducted

Table A.2: t-ratio tests - Period 1. Correction by DenHaan & Levin. Nota : The dependent variables are in columns, and the independent ones in rows

Second period

Harding and Pagan's test using DenHaan and Levin robust t-ratio

	Euro group											EU without Euro			Externals			
	Au	Be	Fi	Fr	Ge	Gr	It	Lux	Net	Po	Sp	De	Swe	UK	No	Swi	Ja	US
Au		*	-	*	**	**	-	-	**	-	-	-	-	-	-	-	-	-
Be	*		-	-	-	*	-	-	-	-	-	-	-	-	-	-	-	-
Fi	-	-		-	-	*	-	-	-	-	-	-	*	-	-	-	-	-
Fr	**	-	-		***	**	-	*	**	-	-	-	-	-	-	-	-	-
Ge	**	-	-	**		**	-	-	**	-	-	-	-	-	-	-	-	-
Gr	**	-	-	-	*		-	-	*	-	-	-	-	-	-	-	-	-
It	-	-	-	-	-	*		-	-	-	-	-	-	-	*	-	-	-
Lux	-	-	-	*	-	-	-		-	*	-	-	-	**	-	-	-	-
Net	**	-	-	**	**	*	-	-		-	-	-	-	-	*	-	-	-
Po	-	-	-	-	-	-	-	-	-		-	-	-	-	-	-	-	-
Sp	-	-	-	-	-	-	-	-	-	-		-	-	-	-	*	-	-
De	-	-	-	-	-	-	-	-	-	-	-		-	-	-	-	-	-
Swe	-	-	-	-	-	-	-	-	-	-	-	-		-	-	-	-	-
UK	-	-	-	-	-	-	-	-	-	-	-	-	-		-	-	-	-
No	-	-	-	-	-	-	-	-	*	-	-	-	-	-		-	-	-
Swi	-	-	-	-	-	-	-	-	-	-	-	-	-	-	-		-	-
Ja	-	-	-	-	-	-	-	-	-	-	-	-	-	-	-	-		-
US	-	-	-	-	-	-	*	-	-	-	-	-	-	-	**	-	-	

Significance levels: *- : non rejection*

**** : 1%* *Blank: test could not be conducted*

*** : 5%*

** : 10%*

Table A.3: t-ratio tests - Period 2. Correction by DenHaan & Levin

First period

Harding and Pagan's test using quasi-difference correction

	Euro group											EU without Euro			Externers			
	Au	Be	Fi	Fr	Ge	Gr	It	Lux	Net	Po	Sp	De	Swe	UK	No	Swi	Ja	US
Au		***	***	***	***	**	***	***	***	***	***		***	***	***	***	***	***
Be	***		***	***	***	***	***	***	***	***	***	*	***	***	***	***	***	***
Fi	***	***		***	***	-	***	***	***	***	***		***	***	***	***	***	***
Fr	***	***	***		***	-	***	***	***	***	***	***	***	***	***	***	***	***
Ge	***	***	***	***		***	***	***	***	***	***		*	***	***	-	***	***
Gr	***	***	-	*	***		-	-	*	***	-		*	*	-	*	***	***
It	***	***	***	***	***	-		-	***	***	***	-	***	***	***	***	***	***
Lux	***	***	***	***	***	-	-		***	***	***	***	***	***	**	***	***	***
Net	***	***	***	***	***	-	***	***		***	***		***	***	***	***	***	***
Po	***	***	***	***	***	***	***	***	***		***		***	***	***	*	***	***
Sp	***	***	***	***	***	-	***	***	***	***			***	***	***	-	***	***
De		*		***			-	***					***		*	***		
Swe	***	***	***	***	-	-	***	***	***	***	***	***		***	***	***	***	***
UK	***	***	***	***	***	-	***	***	***	***	***		***		***	-	***	***
No	***	***	***	***	***	-	***	**	***	***	***	*	***	***		***	***	*
Swi	***	***	***	***	*	-	***	***	***	-	*	***	***	*	***		-	*
Ja	***	***	***	***	***	***	***	***	***	***	***		***	***	***	-		***
US	***	***	***	***	***	***	***	***	***	***	***		***	***	*	-	***	

Significance levels:
*** : 1%
** : 5%
* : 10%

- : non rejection
Blank: test could not be conducted

Table A.4: t-ratio tests - Period 1. Correction by quasi-differences

Second period

Harding and Pagan's test using quasi-difference correction

	Euro group											EU without Euro			Externals			
	Au	Be	Fi	Fr	Ge	Gr	It	Lux	Net	Po	Sp	De	Swe	UK	No	Swi	Ja	US
Au		***	***	***	***	***	***	***	***	***	***	-	***	***	***	-	***	***
Be	***		-	***	***	***	***	-	***	*	**	-	-	-	-	-	***	***
Fi	***	-		***	***	***	***	***	-	***	***	-	***	***	***	-	**	***
Fr	***	***	***		***	***	***	***	***	***	***	***	-	***	***	-	***	***
Ge	***	***	***	***		***	***	***	***	***	***	***	-	***	***	-	***	***
Gr	***	***	***	***	***		***	***	***	***	***	-	***	***	***	-	***	***
It	***	***	***	***	***	***		***	***	-	***	-	-	*	***	***	-	***
Lux	***	-	***	***	***	***	***		***	***	***	***	-	***	***	***	***	-
Net	***	***	-	***	***	***	***	***		***	-	***	***	-	***	-	***	***
Po	***	-	***	***	***	***	-	***	***		-	***	***	-	***	-	***	-
Sp	***	**	***	***	***	***	***	***	-	-		***	-	***	-	***	***	***
De	-	*	-	***	***	*	-	***	***	***	***		***	-	-	***	*	-
Swe	**	-	***	-	-	***	-	-	***	***	-	***		-	***	-	***	-
UK	***	-	***	***	***	***	-	***	-	-	***	-	-		-	***	***	-
No	***	-	***	***	***	***	***	***	***	***	-	-	***	-		-	-	***
Swi	-	-	-	-	-	-	***	***	-	-	***	***	-	***	-		***	-
Ja	***	***	*	***	***	***	-	***	***	***	***	**	***	***	-	***		-
US	***	***	***	***	***	***	***	-	***	-	***	-	-	-	***	-	-	

Significance levels:
**** : 1%*
*** : 5%*
** : 10%*

- : non rejection
Blank: test could not be conducted

Table A.5: t-ratio tests - Period 2. Correction by quasi-differences

A.2 Appendix for Chapter 2

A.2.1 Additional tables for 'distinguishing common and transmitted cycles'

Gruben et al. (2002) OLS estimations

Regressions of comovements on IIT - GKM approach

	Dummies none	time	EU group	time & EU
Dept. variables				
- Coefficients				
Corrdiff	0.16 ***	0.18 ***	0.16 ***	0.18 ***
Corrbk	0.51 **	0.62 ***	0.49 ***	0.59 ***
Coher	0.46 ***	0.51 ***	0.36 ***	0.43 ***
Coherdiff	0.18 ***	0.30 ***	0.18 ***	0.27 ***
Coherbk	0.32 ***	0.33 ***	0.30 ***	0.33 ***
- Wald tests				
Corrdiff	NA			***
Corrbk	NA	***	*	***
Coher	NA	***	***	***
Coherdiff	NA	***		***
Coherbk	NA	***		***

Regressions of comovements on weighted IIT - GKM approach

	Dummies none	time	EU group	time & EU
Dept. variables				
- Coefficients				
Corrdiff	3.12 **	3.51 **	3.42 **	3.53 **
Corrbk	8.76 ***	9.86 ***	7.92 ***	8.86 ***
Coher	11.27 ***	11.87 ***	7.48 ***	8.37 ***
Coherdiff	4.83 ***	6.79 ***	4.90 ***	5.84 ***
Coherbk	4.50 **	4.12 **	3.65 *	4.11 **
- Wald tests				
Corrdiff	NA			***
Corrbk	NA	***		***
Coher	NA	***	***	***
Coherdiff	NA	***	*	***
Coherbk	NA	***	***	***

Regressions of phase lags on IIT - GKM approach

	Dummies none	time	EU group	time & EU
Dept. variables				
- Coefficients				
Aphlag	-0.4 *	-0.50 **	-0.06	-0.18
Aphlagdiff	-1.19 ***	-1.42 ***	-1.04 ***	-1.27 ***
Aphlagbk	-1.22 ***	-1.52 ***	-1.24 ***	-1.50 ***
- Wald tests				
Aphlag	NA	***	***	***
Aphagdiff	NA	***	***	***
Aphlagbk	NA	***	***	***

Regressions of phase lags on weighted IIT - GKM approach

	Dummies none	time	EU group	time & EU
Dept. variables				
- Coefficients				
Aphlag	-17.4 **	-19.89 ***	-5.43	-6.50
Aphlagdiff	-17.94 ***	-20.43 ***	-11.04 *	-13.36 **
Aphlagbk	-14.36 *	-17.57 ***	-13.55 **	-15.81 ***
- Wald tests				
Aphlag	NA	***	***	***
Aphlagdiff	NA	***	***	***
Aphlagbk	NA	***		***

Regressions of the Common cycles measure on IIT - GKM approach

	Dummies none	time	EU group	time & EU
Dept. variables				
- Coefficients				
Coms	2.00 ***	2.03 ***	0.89	1.22 **
Comsdiff	2.26 ***	3.09 ***	2.07 ***	2.81 ***
Comsbk	2.91 ***	3.43 ***	2.97 ***	3.43 ***
- Wald tests				
Coms	NA	***	***	***
Comsdiff	NA	***	***	***
Comsbk	NA	***	**	***

Regressions of the Common cycles measure on weighted IIT - GKM

	Dummies none	time	EU group	time & EU
Dept. variables				
- Coefficients				
Coms	57.10 ***	56.00 ***	14.73	20.19
Comsdiff	31.24 **	43.00 ***	21.31	28.97 **
Comsbk	37.05 **	41.96 **	36.10 **	40.59 **
- Wald tests				
Coms	NA	***	***	***
Comsdiff	NA	***		***
Comsbk	NA	***		***

Table A.6:

Regressions of the Transmission measure on IIT - GKM approach

	Dummies none	time	EU group	time & EU
Dept. variables				
- Coefficients				
Trans	-0.03	0.23	0.64	0.68
Transdiff	-0.80 **	-0.75 **	-0.62 **	-0.65 **
Transbk	-1.05 ***	-1.26 ***	-1.02 **	-1.20 ***
- Wald tests				
Trans	NA	***	***	***
Transdiff	NA	**	***	**
Transbk	NA	***	**	***

Regressions of the Transmission measure on weighted IIT - GKM

	Dummies none	time	EU group	time & EU
Dept. variables				
- Coefficients				
Trans	-11.69	-7.50	14.13	13.09
Transdiff	2.40	5.67	13.22	12.72
Transbk	-15.95	-17.82	-13.33	-14.22
- Wald tests				
Trans	NA	***	***	***
Transdiff	NA	**	***	
Transbk	NA	***	***	

Dividing the panel into EU and non-EU countries

Estimated coefficients from the first stage

	Trade Intensity	*IIT*	*weighted IIT*
Whole Panel			
Com. Language	0.40	10.78 ***	0.38 **
Distance	-0.32 ***	-6.97 ***	-0.25 ***
Com. Borders	1.80 ***	9.29 ***	1.09 ***
R-squared	0.32	0.32	0.39
EU countries			
Com. Language	-0.48	5.74	0.06
Distance	-1.44 ***	-17.79 ***	-0.98 ***
Com. Borders	1.91 ***	4.96 *	1.18 ***
R-squared	0.50	0.58	0.54
Non-EU countries			
Com. Language	1.49 ***	18.87 ***	1.12 ***
Distance	0.00	-2.71 ***	-0.02
Com. Borders	1.03 ***	7.71 **	0.56 ***
R-squared	0.34	0.26	0.56

Rows: independent variables
Columns: dependent variables

IV Regressions of phase lags on IIT

	Aphlag	Aphlag diff	Aphlag BK
Whole Panel			
IIT	0.79 **	-0.84 **	-1.67 ***
IIT*Trade intensity	10.61 *	-16.44 ***	-29.28 ***
EU countries			
IIT	0.03	-1.36 ***	-1.53 ***
IIT*Trade intensity	-0.41	-18.01 ***	-22.27 ***
Non-EU countries			
IIT	-0.16	-1.53 *	-2.18 **
IIT*Trade intensity	-9.68 *	-33.00 **	-40.77 **

Rows: independent variables
Columns: dependent variables
Aphlag: absolute value of phase lags.
Aphlag diff: absolute value of phase lags computed on differenced series.
Aphlag: absolute value of phase lags computed on BK-filtered series.

IV Regressions of phase lags on IIT

	Aphlag	Aphlag diff	Aphlag BK
Whole Panel			
IIT	0.79 **	-0.84 **	-1.67 ***
IIT*Trade intensity	10.61 *	-16.44 ***	-29.28 ***
EU countries			
IIT	0.03	-1.36 ***	-1.53 ***
IIT*Trade intensity	-0.41	-18.01 ***	-22.27 ***
Non-EU countries			
IIT	-0.16	-1.53 *	-2.18 **
IIT*Trade intensity	-9.68 *	-33.00 **	-40.77 **

Rows: independent variables
Columns: dependent variables
Aphlag: absolute value of phase lags.
Aphlag diff: absolute value of phase lags computed on differenced series.
Aphlag: absolute value of phase lags computed on BK-filtered series.

IV Regressions of the Transmission measure on IIT

	Trans	Trans diff	Trans BK
Whole Panel			
IIT	1.76 ***	-1.86 ***	-1.67 **
IIT*Trade intensity	25.98 **	-33.75 ***	-28.38 **
EU countries			
IIT	1.82 **	-0.81	-1.24
IIT*Trade intensity	22.95 **	-9.33	-19.28
Non-EU countries			
IIT	-0.33	-0.96	-0.86
IIT*Trade intensity	-17.00	-26.15	-22.23

*Trans: measure of transmission on raw series = log(coher * |phase lags|)*
Trans diff: same as Trans on differenced series
Trans BK: same as Trans on BK-filtered series

IV Regressions of the Common cycles measure on IIT

	Coms	Coms diff	Coms BK
Whole Panel			
IIT	-0.89	2.89 ***	3.41 ***
IIT*Trade intensity	-5.40	56.45 ***	60.69 ***
EU countries			
IIT	-0.22	4.01 ***	2.78 **
IIT*Trade intensity	2.03	54.38 ***	43.77 ***
Non-EU countries			
IIT	2.72	3.49 ***	3.68 **
IIT*Trade intensity	101.13 ***	73.68 ***	64.12 *

Coms: common shocks measured on raw series = log(coher / |phase lags|)
Coms diff: same as Coms on differenced series
Coms BK: same as Coms on BK-filtered series

Regressions of comovements on weighted IIT - GKM approach

	Corr diff	Corr BK	Coher	Coher diff	Coher BK
Whole Panel					
eq.1: IIT	0.16 ***	0.51 ***	0.46 ***	0.18 ***	0.32 ***
eq.2: IIT*TI	3.12 **	8.76 ***	11.27 ***	4.83 ***	4.50 **
(1-IIT)*TI	2.36	0.40	-5.03 *	-0.78	0.67
EU countries					
eq.1: IIT	0.35 ***	0.74 ***	0.82 ***	0.27 ***	0.37 *
eq.2: IIT*TI	11.18 ***	10.66 *	6.37	4.22	11.45 *
(1-IIT)*TI	-11.46 **	-6.33	-3.10	-3.73	-14.63 *
Non-EU countries					
eq.1: IIT	0.22 **	0.18	0.56 ***	0.03	0.34 ***
eq.2: IIT*TI	3.08	4.72	8.87 *	5.30	10.77 ***
(1-IIT)*TI	3.66 **	3.19	-2.83	-0.12	0.80

TI: 'Trade intensity'

Regressions of phase lags on IIT - GKM approach

	Aphlag	Aphlag diff	Aphlag BK
Whole Panel			
eq.1: IIT	-0.40 *	-1.19 ***	-1.22 ***
eq.2: IIT*TI	-17.40 **	-17.94 ***	-14.36 *
(1-IIT)*TI	1.89	1.41	-1.96
EU countries			
eq.1: IIT	-0.27	-0.68	-1.65 ***
eq.2: IIT*TI	-0.12	-14.78	-28.89 *
(1-IIT)*TI	-2.16	10.18	26.66
Non-EU countries			
eq.1: IIT	-0.20	-0.08	-0.76
eq.2: IIT*TI	-15.95	-28.31 *	-21.72
(1-IIT)*TI	-3.50	1.30	-2.95

TI: 'Trade intensity'

Regressions of the Transmission measure on IIT - GKM

	Trans	Trans diff	Trans BK
Whole Panel			
eq.1: IIT	-0.03	-0.80 **	-1.05 **
eq.2: IIT*Trade intensity	-11.69	2.40	-15.95
(1-IIT)*Trade intensity	9.12	-5.69	9.62
EU countries			
eq.1: IIT	1.21	0.67	-1.88
eq.2: IIT*Trade intensity	22.87	26.75	-29.86
(1-IIT)*Trade intensity	0.53	-18.47	38.99
Non-EU countries			
eq.1: IIT	0.05	0.30	-0.72
eq.2: IIT*Trade intensity	-45.29	11.74	25.66
(1-IIT)*Trade intensity	15.73	14.97	-17.20

*Trans: measure of transmission on raw series = log(coher * |phase lags|)*
Trans diff: same as Trans on differenced series
Trans BK: same as Trans on BK-filtered series
TI: 'Trade intensity'

Regressions of the Common cycles measure on IIT - GKM

	Coms	Coms diff	Coms BK
Whole Panel			
eq.1: IIT	2.00 **	2.26 ***	2.91 ***
eq.2: IIT*Trade intensity	57.10 ***	31.24 **	37.05 **
(1-IIT)*Trade intensity	-19.31	2.94	-0.05
EU countries			
eq.1: IIT	2.40 *	2.12 **	4.46 ***
eq.2: IIT*Trade intensity	5.56	7.93	93.07 **
(1-IIT)*Trade intensity	-3.53	-8.65	-111.13 **
Non-EU countries			
eq.1: IIT	2.59 **	0.11	3.44 **
eq.2: IIT*Trade intensity	86.33	50.96	83.25
(1-IIT)*Trade intensity	-14.03	-17.04	-20.36

Coms: common shocks measured on raw series = log(coher / |phase lags|)
Coms diff: same as Coms on differenced series
Coms BK: same as Coms on BK-filtered series
TI: 'Trade intensity'

A.2.2 Additional tables for 'distinguishing vertical and horizontal IIT'

First stage estimations, pooled OLS estimations

	Horizontal IIT		Vertical IIT		One-way trade	
Borders	0.108 ***	0.107 ***	0.004	-0.049	-0.093 ***	-0.016 **
Language	0.053 ***	0.059 ***	-0.013	0.025	0.006	-0.075 ***
Log(distance)	0.116 ***	0.118 ***	0.042	-0.032	-0.085 ***	0.022 **
Market size		-1.72E-09		7.70E-03 ***		-1.21E-07 ***
Dem. for differ.		-8.25E-06 **		3.56E-06		8.98E-06 ***
Comp. adv.		1.35E-05 ***		-5.34E-05 ***		1.51E-05 **
R2	0.44	0.53	0.04	0.25	0.27	0.54

First stage estimations, fixed time effect

	Horizontal IIT		Vertical IIT		One-way trade	
Borders	0.107 ***	0.107 ***	0.005	-0.050	-0.093 ***	-0.015 *
Language	0.056 ***	0.059 ***	-0.015	0.003	0.004	-0.068 ***
Log(distance)	0.119 ***	0.118 ***	0.040 **	-0.046	-0.086 ***	0.027 *
Market size		3.12E-09		2.19E-08		-1.02E-07 ***
Dem. for differ.		1.25E-07		-7.84E-05 ***		3.74E-05 ***
Comp. adv.		9.32E-06		-5.34E-05 ***		1.62E-05 **
R2	0.58	0.59	0.12	0.4	0.31	0.58

First stage estimations, group effect

	Horizontal IIT		Vertical IIT		One-way trade	
Borders	0.007	0.158 ***	0.133836 ***	-0.046	0.039 ***	-0.061 ***
Language	0.061 ***	0 ??	-0.0667463 ***	0 ??	0.013	0 ??
Log(distance)	0.039 ***	0.046 ***	0.0913475 ***	0.103 ***	0.015 ***	0.015 ***
Market size		-1.36E-07 *		1.79E-07 *		1.76E-08
Dem. for differ.		-1.41E-06		-3.50E-06		3.65E-06
Comp. adv.		2.01E-06		-1.15E-05 *		-3.05E-06
R2	0.54	0.63	0.63	0.67	0.82	0.85

First stage estimations, within transformation

	Horizontal IIT		Vertical IIT		One-way trade	
Borders	0.104 ***	0.103 ***	0.007	-0.053	-0.092 ***	-0.011
Language	0.061 ***	0.065 ***	-0.007	0.030	3.14E-04	-0.088 ***
Log(distance)	0.123 ***	0.124 ***	0.049 **	-0.051	-0.090 ***	0.028 *
Market size		4.45E-09		6.32E-08 **		-1.24E-07 ***
Dem. for differ.		3.86E-06		-4.94E-05 *		2.86E-05 ***
Comp. adv.		5.73E-06		-5.84E-05 ***		1.94E-05 ***
R2	0.54	0.55	0.05	0.34	0.29	0.61

First stage estimations, within transformation and time effect

	Horizontal IIT		Vertical IIT		One-way trade	
Borders	0.110 ***	0.108 ***	-0.019	-0.067 *	-0.089 ***	-0.013
Language	0.056 ***	0.060 ***	0.018 *	0.042	-0.003	-0.083 ***
Log(distance)	0.120 ***	0.116 ***	0.066 ***	-0.029	-0.092 ***	0.036 **
Market size		2.93E-09		5.07E-08		-1.18E-07 ***
Dem. for differ.		-3.92E-06		-5.11E-05 *		4.16E-05 ***
Comp. adv.		6.77E-06		-5.99E-05 ***		2.05E-05 **
R2	0.63	0.64	0.22	0.49	0.31	0.64

A.3 Appendix for Chapter 3

A.3.1 Model

Baseline model

We use the structural decomposition of the Harvey type (Harvey, 1989). Each observed variable i is determined by two unobserved components, a trend $\mu_{i,t}$ and a cycle $\psi_{i,t}$.

$$y_{i,t} = \alpha_i^{'}\mathbf{x}_{i,t} + \mu_{i,t} + \psi_{i,t} + \varepsilon_{i,t}, t = 1, ..., T, i = 1, ..., k \tag{A.11}$$

$\mathbf{x}_{i,t}$ corresponds to external variables –seasonal dummies and information variables for outliers and structural breaks. See the description of the deJong & Penzer procedure below. α_i denotes corresponding coefficients. Note that in order to ease the presentation, we drop this term below.

Suppose also that $\varepsilon_{i,t} \sim NID(0, \sigma_{\varepsilon i}^2)$. In addition, suppose that the trend has the 'local linear' form –i.e. is a random walk with a drift that is itself a random walk.

$$\mu_{i,t} = \mu_{i,t-1} + \beta_{i,t-1} + u_{i,t},\ u_{i,t} \sim NID(0, \sigma_{i,u}^2) \tag{A.12}$$

$$\beta_{i,t} = \beta_{i,t-1} + v_{i,t},\ v_{i,t} \sim NID(0, \sigma_{i,v}^2) \tag{A.13}$$

The cycle $\psi_{i,t}$ is part of the process generated by

$$\begin{pmatrix} \psi_{i,t} \\ \psi_{i,t}^{*} \end{pmatrix} = \rho_i \begin{pmatrix} \cos\omega_i & \sin\omega_i \\ -\sin\omega_i & \cos\omega_i \end{pmatrix} \begin{pmatrix} \psi_{i,t-1} \\ \psi_{i,t-1}^{*} \end{pmatrix} + \begin{pmatrix} \kappa_{i,t} \\ \kappa_{i,t}^{*} \end{pmatrix} \tag{A.14}$$

ρ is the so-called 'damping factor'. It belongs to the interval [0,1]. Note that (A.14) is simply a way to put in a state space form the oscillating process $\rho_i cos\omega_i t + \rho_i sin\omega_i t + \kappa_{i,t}$ Besides, as shown by Harvey (1989), the cyclical process can be rewritten as an ARMA(2,1) and $0 \leq \rho \leq 1$ is a requirement for the roots to lie outside the unit circle. At the limit, when $\rho = 0$, the cycle becomes purely white noise. When $\rho = 1$, the process exhibits a unit root. The error terms $\kappa_{i,t}$ and $\kappa_{i,t}^{*}$ are supposed to be *NID* with mean zero and variances $\sigma_{i,\kappa}^2$ and σ_{i,κ^*}^2, respectively.

Note that it is also assumed that the error terms are uncorrelated, such that

$$\Omega_i = \begin{pmatrix} \sigma_{\varepsilon_i}^2 & 0 \\ 0 & \mathbf{P}_i \end{pmatrix} \text{ where } \mathbf{P}_i = diag(\ \sigma_{i,u}^2 \quad \sigma_{i,v}^2 \quad \sigma_{i,\kappa}^2 \quad \sigma_{i,\kappa*}^2\ ) \tag{A.15}$$

We will see below that a restriction is imposed in that $\sigma_{i,\kappa}^2 = \sigma_{i,\kappa^*}^2$.

Putting this univariate model into a state space form gives,

$$y_{i,t} = \mathbf{z}\zeta_{i,t} + \varepsilon_{i,t} \tag{A.16}$$

$$\zeta_{i,t+1} = \tau_i\zeta_{i,t} + \eta_{i,t} \tag{A.17}$$

with $\zeta_{i,t}' = \begin{pmatrix} \mu_{i,t} & \beta_{i,t} & \psi_{i,t} & \psi_{i,t}^* \end{pmatrix}$ and $\eta_{i,t}' = \begin{pmatrix} u_{i,t} & v_{i,t} & \kappa_{i,t} & \kappa_{i,t}^* \end{pmatrix}$. Besides, $\mathbf{z} = \begin{pmatrix} 1 & 0 & 1 & 0 \end{pmatrix}$ and τ_i is a (4×4) matrix:

$$\tau_i = \begin{pmatrix} 1 & 1 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & \alpha_i & \beta_i \\ 0 & 0 & \gamma_i & \delta_i \end{pmatrix}$$

where the sub-matrix $\begin{pmatrix} \alpha_i & \beta_i \\ \gamma_i & \delta_i \end{pmatrix}$ is the coefficient matrix of the cyclical component defined in (A.14). (A.16) is the measurement equation and (A.17) is the state equation. $\mathbf{z}\zeta_{i,t} = y_{i,t} - \varepsilon_{i,t}$ is the signal.

Multivariate setting

Leaving apart any common component consideration, the multivariate specification of the model is straightforward. We get

$$\underset{k\times 1}{\mathbf{y}_t} = \mathbf{Z}\zeta_t + \varepsilon_t \tag{A.18}$$

$$\underset{km\times 1}{\zeta_{t+1}} = \mathbf{T}\zeta_t + \eta_t \tag{A.19}$$

with y_t being a vector of observed variables, $\zeta_t' = \begin{pmatrix} \zeta_{1,t}' & ... & \zeta_{k,t}' \end{pmatrix}$ a vector of unobserved components. $\mathbf{Z} = I_k \odot \mathbf{z}$ and $\mathbf{T} = diag\begin{pmatrix} \tau_1 & ... & \tau_k \end{pmatrix}$ are coefficients matrices. m is the number of state variables and k is the number of countries. In the present case, m= 4. However, this is no longer the case when additional state variables are added to the model[17] .

[17]Namely, seasonal dummies and information variables for outliers and structural breaks. Typically, the number of information variables will vary from one country to the other, depending on the tests implemented -see below- so that m is not known in advance.

Common components

Assume now that each observed series is not only determined by the idiosyncratic unobserved variables of (A.11), but also by components common to all the other series. Common elements are treated as principal factors. The task of this paper is much more modest than in the seminal article of Forni et. al. (2000). First because we have far less variables and second because we make the assumption that there is just one common cycle. This assumption obliges us to include idiosyncratic elements as well for each of the variables. Otherwise, this would come down to assuming that every element of the dependent vector is entirely determined by the common element, while the remaining part is white noise. This would certainly be a too strong assumption. Thus there are two common factors, one common trend and one common cycle. The assumption that there is only one common factor of each type is relatively strong, but simplifies the interpretation in terms of a *European cycle* or any other common cycle, and is necessary for our task.

$$y_{i,t} = \mu_{i,t} + \psi_{i,t} + \theta_i \widetilde{\mu}_t + \omega_i \widetilde{\psi}_t + \varepsilon_{i,t}, t = 1, ..., T, i = 1, ..., k \tag{A.20}$$

where μ_t and ψ_t are the common trend and common cycle, respectively, and τ_i and ω_i the corresponding loading coefficients for country i. The assumption is made that both (absolute values of the) coefficients belong to the interval [0,1], which is equivalent to assuming that the common elements do not determine the series as much as the idiosyncratic part. The multivariate model becomes

$$\underset{k\times 1}{\mathbf{y}_t} = \mathbf{Z}^* \zeta_t^* + \varepsilon_t \tag{A.21}$$

$$\underset{(mk+m)\times 1}{\zeta_{t+1}^*} = \mathbf{T}^* \zeta_t^* + \eta_t \tag{A.22}$$

where

$$Z^* = \begin{pmatrix} Z & \Theta & 0_{k\times 1} & \Omega & 0_{k\times 1} \end{pmatrix}$$

with Θ and Ω being $k \times 1$ vectors containing the loading coefficients and $\mathbf{0}_{k\times 1}$ being vectors of zeros. Also

$$\underset{(4k+4)\times(4k+4)}{\mathbf{T}^*} = diag(\ \mathbf{T} \quad \tau \ ) = diag \begin{pmatrix} \tau_1 & \dots & \tau_k & \tau \end{pmatrix}$$

$$\zeta_t^{*\prime} = (\ \zeta_t' \quad \zeta_t' \ ) = \begin{pmatrix} \zeta_{1,t}' & \dots & \zeta_{k,t}' & \zeta_t' \end{pmatrix}$$

Estimation

The idiosyncratic and common cycles are estimated by the Kalman smoother. A presentation of the Kalman filter and smoother goes well beyond the scope of this paper. See Harvey (1989 chap.3) for an extensive treatment of the topic. See also Proietti (2002, appendix C). All computations were performed with the Ox library SsfPack version 2.2 of Koopman, Shephard and Doornik[18]. For a documentation on this library, see Koopman et al. (1999). Codes and data are available upon request.

Maximum Likelihood estimation and identifiability The parameters are estimated first by maximum likelihood. The procedure used here for MLE is the one of Koopman et al. (1999, p.140), based upon the BFGS numerical optimisation method. Some restrictions must be imposed on the parameters of the model in order to get full identification. We impose first positive definiteness of the variances. Harvey (1989) shows that the local linear trend model is identified, provided that the disturbances are normally distributed and mutually uncorrelated. He shows also that the conditions $\rho > 0$, $\omega \in [0, \pi]$ and $E[\kappa_t \kappa_t^*] = 0$ in the stochastic cycle model are sufficient to insure identifiability. Note that $\sigma^2_{i,\kappa} = \sigma^2_{i,\kappa*}$ is not required. However, this restriction has been imposed on the model as it is a common practice in the literature (e.g. Harvey & Jaeger, 1993 or Proietti, 2002) and as estimations seemed more stable with this additional restriction. To sum up, the main restrictions are given by (A.15), ρ and ω ($\rho \in [0, 1]$ and $\omega \in [0, \pi]$, $\forall i$)

Additional restrictions are placed upon the common components. It is common in dynamic factor analysis to set the error terms covariance matrix of the common factors equal to an identity matrix. However, since we do want to compare the share of variances due to the common cycle, the error term variance of the common part was set equal to the mean of the error terms of the idiosyncratic parts, such that $\sigma_e^2 = \frac{1}{k} \sum\limits_i \sigma^2_{i,e}$, $e = \{u, v, \kappa\}$. For some reasons, ρ, the damping factor of the common cycle, tends to increase with the number of countries in the group under consideration. To avoid this problem, we set $\rho = \frac{1}{k} \sum\limits_i \rho_i$.

Setting starting values for the Kalman filter The Kalman filter estimates optimally the value of the state vector at time t. But this optimal estimation requires that the

[18] available at www.ssfpack.com

initial state vector ζ_0^* and the covariance matrix of the initial state vector, $\mathbf{P}_0$, are known. Usually, ζ_0^* and $\mathbf{P}_0$ are set such that they are the mean and covariance matrix of the unconditional distribution of ζ_t^*. Harvey (1989, p.121) shows that for the stochastic cycle model, the initial conditions are a zero mean and a covariance matrix $\mathbf{P}_0 = \left[\sigma_\kappa^2/(1-\rho^2)\right] \mathbf{I}_2$, where σ_κ^2 and ρ^2 are the same as above and I_2 is a (2×2) identity matrix. Given the values of σ_κ^2 and ρ^2 found with the ML estimations, $\mathbf{P}_0$ would oscillate between 10^{-2} and 10^{-3}. In order to simplify the calculations, we have taken an arbitrarily small value and set $\mathbf{P}_0 = 10^{-8}$ here. The results should not be too much affected by this choice.

When the state equation is not stationary, which is the case for the local linear trend, the unconditional distribution of the state vector is not available. One must therefore use a non-informative initial condition. This is best achieved by using a diffuse prior such that $\mathbf{P}_0 = v\mathbf{I}$, where $v \to \infty$. In practice, we follow Koopman et al. (1999) who use $v = 10^6$

A.3.2 Rolling maximum correlations

- Effects of including the UK into the Euro group: 'within' correlations

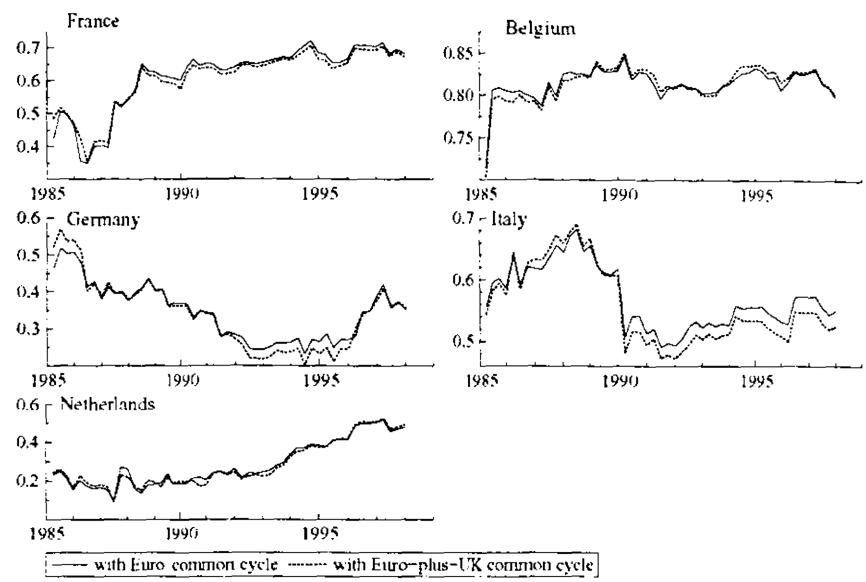

Figure A.18: Maximum rolling correlations of individual cycles with Euro or 'Euro-plus-UK' common cycles – OUTPUT

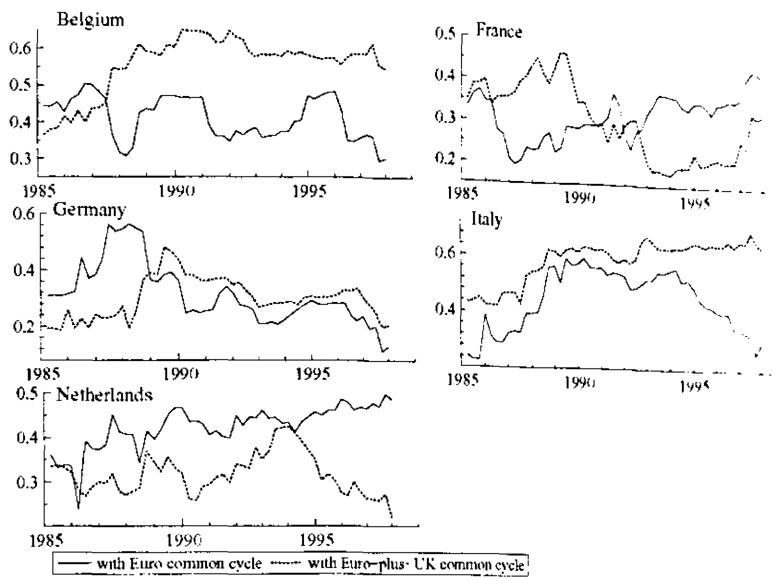

Figure A.19: Maximum rolling correlations of individual cycles with Euro or 'Euro-plus-UK' common cycles – CONSUMPTION

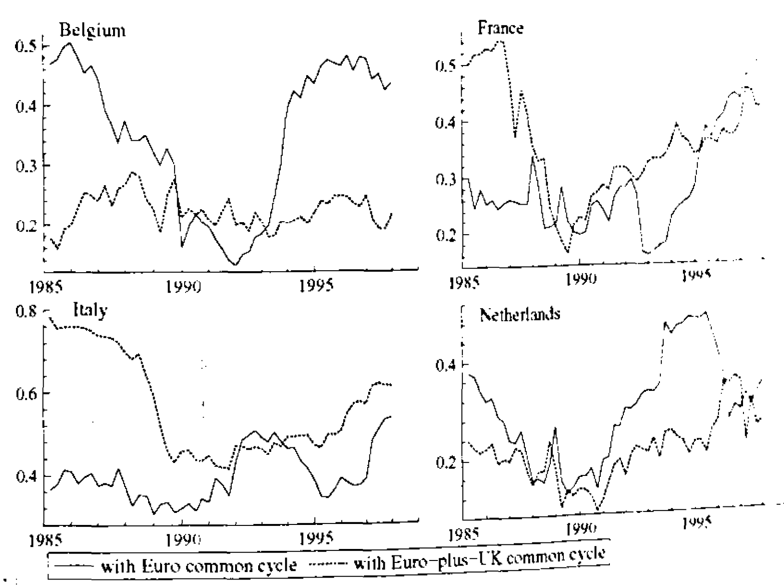

Figure A.20: Maximum rolling correlations of individual cycles with Euro or 'Euro-plus-UK' common cycles – PUBLIC EXPENDITURES

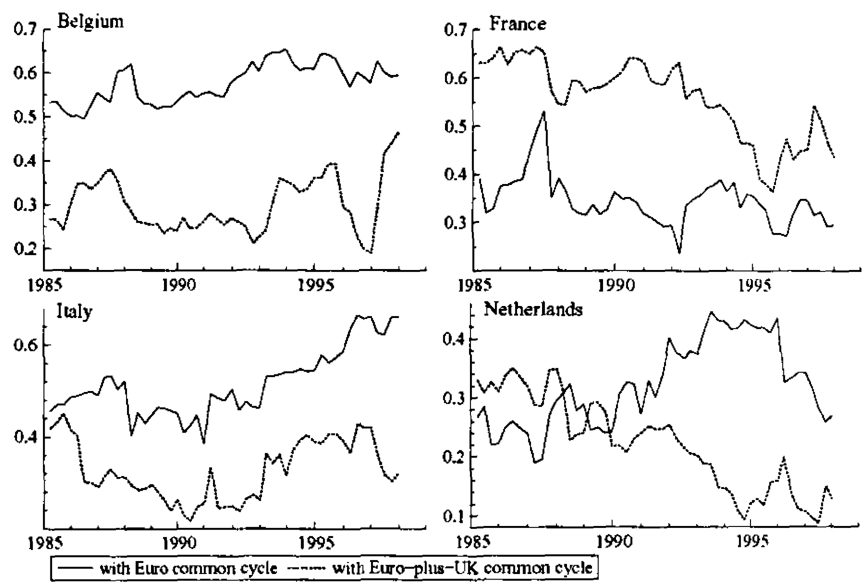

Figure A.21: Maximum rolling correlations of individual cycles with Euro or 'Euro-plus-UK' common cycles – INVESTMENT

A.3.3 Algorithm used

1. Put model into state space form.

2. Estimate the parameters by maximum likelihood (ML).

3. Compute de Jong and Penzer outliers and structural break tests and create dummies accordingly.

4. Incorporate dummies into the state space model.

5. Re-estimate the parameters by ML.

6. Compute the Kalman smoother and extract relevant smoothed state variables (here idiosyncratic and common cycles).

A.3.4 de Jong & Penzer (1998) outliers and structural breaks tests

An important question which is often underestimated in the empirical business cycles literature, is the one of outliers and structural breaks. Indeed, abnormal movements in the data due to measurement errors or extreme shocks affecting the economy should be excluded from the analysis. If this is not the case, the estimated model may be biased.

Tests for structural breaks and outliers in a state-space framework[19] were presented by de Jong and Penzer (1998). See also Koopman et al. (1999). Thanks to the flexibility of the SsfPack library, the implementation of such tests is relatively simple. The test is based on the residuals associated with the state space model, estimated via the Kalman filter. Harvey and Koopman (1992) show that these so-called auxiliary residuals can be used to detect outliers and structural breaks.

The idea behind the test is that the system is driven by some unobserved components (UC, or state variables) which capture its true behaviour. Recall that the Kalman filter optimally extracts state variables. If the model is chosen adequately, the observed series –or measurement equation– is then the sum of a signal and a white noise. However, some shocks to the system cannot be taken into account properly by the model, i.e. outliers and structural breaks. If the disturbance term exhibit aberrant behaviour at some point and

[19]Note that the tests proposed by the authors can be seen more generally as tests for unusual behaviour of the state variables. One could test as well for abnormal cycles or even abnormal moving average terms –in the VAR state-space framework.

cannot be considered as white noise, there must be such a shock at this point. Of course, these tests work under the assumption that the model is the true one, i.e. that the state equations are chosen adequately.

Once the tests have been conducted and that the outliers and structural breaks have been detected, the corresponding information vectors are added to the model. For outliers, they consist in dummy variables equal to one at the time of the outlier and zero elsewhere. For structural breaks, the variable is equal to zero before the date of the break and equal to one after.

Consider the state space system composed of eq.(A.18) and (A.19) and modify the notation slightly such that:

$$\underset{k\times 1}{\mathbf{y}_t} = \mathbf{Z}\zeta_t + \mathbf{G}_t\varepsilon_t \tag{A.23}$$

$$\underset{m\times 1}{\zeta_{t+1}} = \mathbf{T}\zeta_t + \mathbf{H}_t\varepsilon_t \tag{A.24}$$

Suppose that there is an outlier or a level shift at time t. The traditional way to test for this is to include an explanatory variable that takes such a shock into account – e.g. a dummy equal to one at the time of the sock and zero elsewhere. It is sufficient to regress this variable onto the dependent variables and to test if the coefficient is different from zero or not. Let δ be this coefficient, D an explanatory variable and $\sigma^2\Sigma$ the covariance matrix of y_t. Using GLS, we get

$$\widehat{\delta} = (D'\Sigma^{-1}D)^{-1}D'\Sigma^{-1}y = S^{-1}s$$

In order to test $H_0 : \delta = 0$ the following statistic might be used

$$\tau^2 = \widehat{\sigma}^2 s' S^{-1} s$$

which has an approximate χ^2_p distribution where p is the number of linearly independent columns of D. de Jong & Penzer show[20] that the maximum ρ^{*2} of $\rho^2 = \tau^2/\widehat{\sigma}^2$ is

$$\rho_t^{*2} = \mathbf{v}_t'\mathbf{F}_t^{-1}\mathbf{v}_t + \mathbf{r}_t'\mathbf{N}_t^{-1}\mathbf{r}_t \tag{A.25}$$

where $\mathbf{v}_t = \mathbf{G}_t\varepsilon_t$ is the prediction error and F_t is the prediction variance of y_t. v_t and F_t are estimated by the Kalman filter. r_t is the backward prediction error and N_t is the backward prediction variance of ζ_t. r_t and N_t are estimated by the Kalman smoother. Therefore, the first term in the RHS of (A.25) corresponds to the observation equation

[20] theorem 3, p 801.

and the second term corresponds to the state equation. It is not difficult to see from (A.23) and (A.24) that the former indicates abnormal shocks that affect y_t for one period only (i.e. outliers) whereas the latter indicates shocks that affect the level of y_t permanently (i.e. structural breaks). Two important results of de Jong & Penzer are that the maximum (A.25) is also attained when the shocks are uncorrelated and that $\rho_t^{*2} = \mathbf{r}_t'\mathbf{N}_t^{-1}\mathbf{r}_t$ when there is no outlier and $\rho_t^{*2} = \mathbf{u}_t'\mathbf{M}_t^{-1}\mathbf{u}_t$ when there is no shock to the trend. u_t is the backward prediction error and M_t is the backward prediction variance of y_t, obtained from the Kalman smoother. This implies that it is possible to test separately for the different types of abnormal shocks. The authors note that both terms of the RHS of (A.25) follow chi-squared distributions with degrees of freedom equal to the number of components in the measurement and the state equation, respectively. Therefore, ρ_t^{*2} follows a chi-squared distribution as well.

A.3.5 Decomposition of S_i

Common cycle variance to idiosyncratic cycle variance ratio

	Output			Consumption			Public Expenditures			GFCF		
	Euro	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*
Belgium	0.29	0.32		0.23	0.13		0.13	0.25		0.58	0.57	
France	8.94	6.15		0.83	0.57		1.84	6.22		12.93	21.19	
Germany	0.58	0.91		0.78	0.48							
Italy	0.89	0.64		0.09	6.20		7.42	42.69		1.19	2.25	
Netherlands	1.72	1.89		0.24	0.15		0.37	0.72		0.10	0.11	
UK		2.53	0.62		1.70	0.79		0.41	0.45		82.34	1.28
Japan			0.42			0.13			1.72			0.04
US			0.92			5.23			0.67			0.05
Average	2.49	2.07	0.65	0.43	1.54	2.05	2.44	10.06	0.95	3.70	21.29	0.46

Loading coefficients

	Output			Consumption			Public Expenditures			GFCF		
	Euro	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*	*Euro*	*Euro-UK*	*UK-Jp-US*
Belgium	0.79	0.80		-0.61	0.31		0.11	0.38		0.40	-0.12	
France	0.75	0.78		0.15	-0.24		0.22	-0.53		-0.10	0.23	
Germany	0.67	0.70		0.05	-0.04							
Italy	0.67	0.65		-0.11	0.40		-0.44	0.49		0.37	-0.58	
Netherlands	-0.05	-0.03		0.69	-0.55		0.17	-0.08		-0.30	0.22	
UK		0.34	0.49		0.59	0.18		0.54	0.50		0.48	0.39
Japan			0.38			0.40			0.51			0.26
US			0.37			0.44			0.56			-0.05

A.3.6 Spectral densities

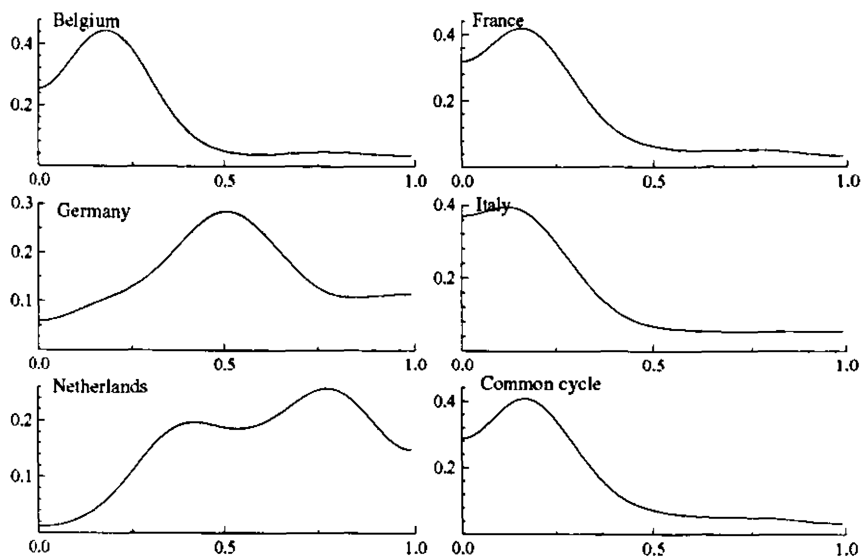

Figure A.22: Spectral densities for the Euro group – Output

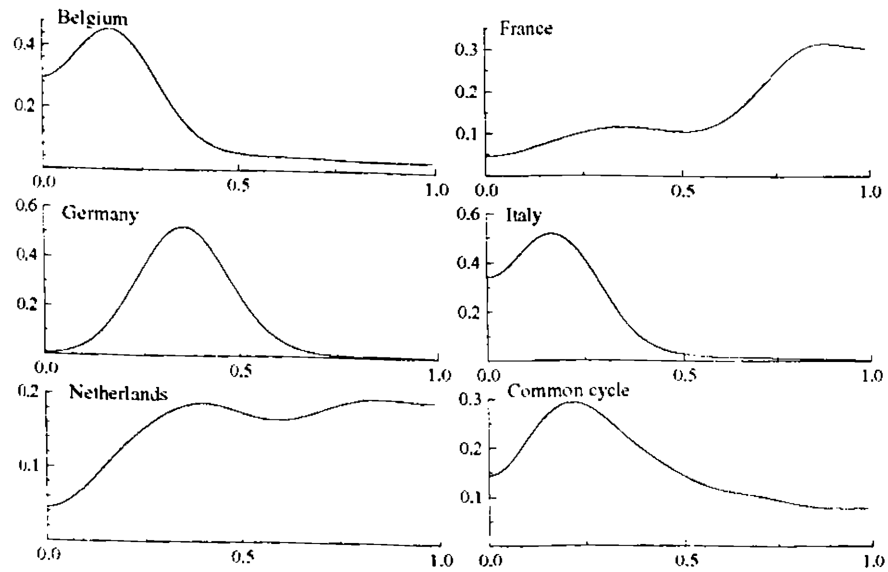

Figure A.23: Spectral densities for the Euro group – Consumption

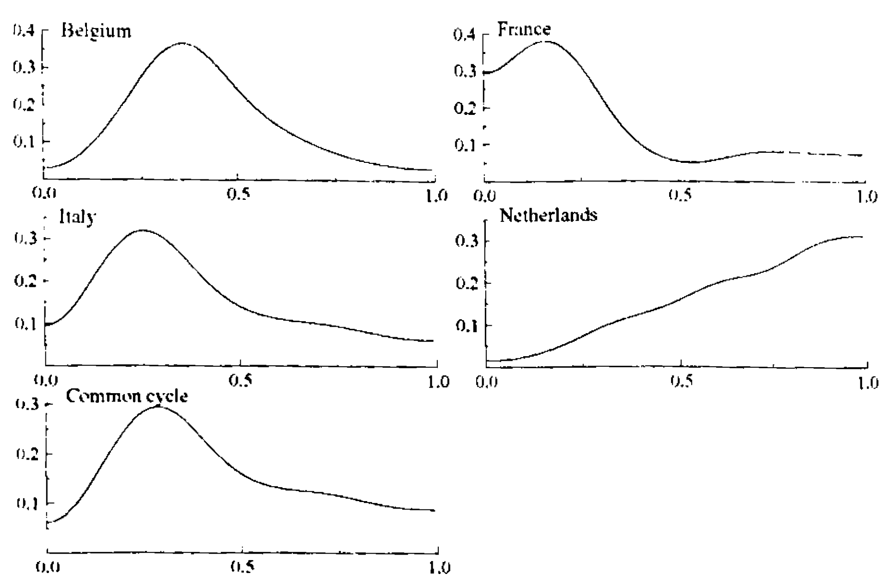

Figure A.24: Spectral densities for the Euro group – Public expenditures

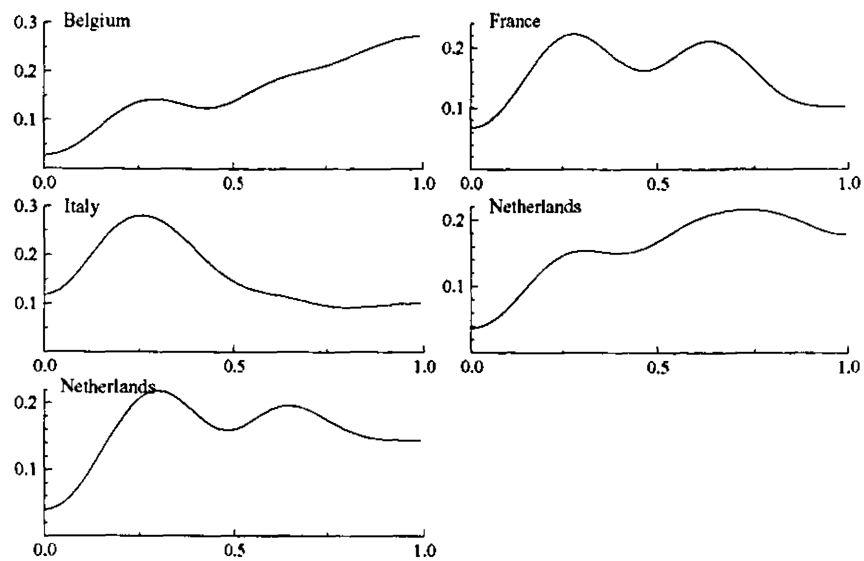

Figure A.25: Spectral densities for the Euro group – Investment

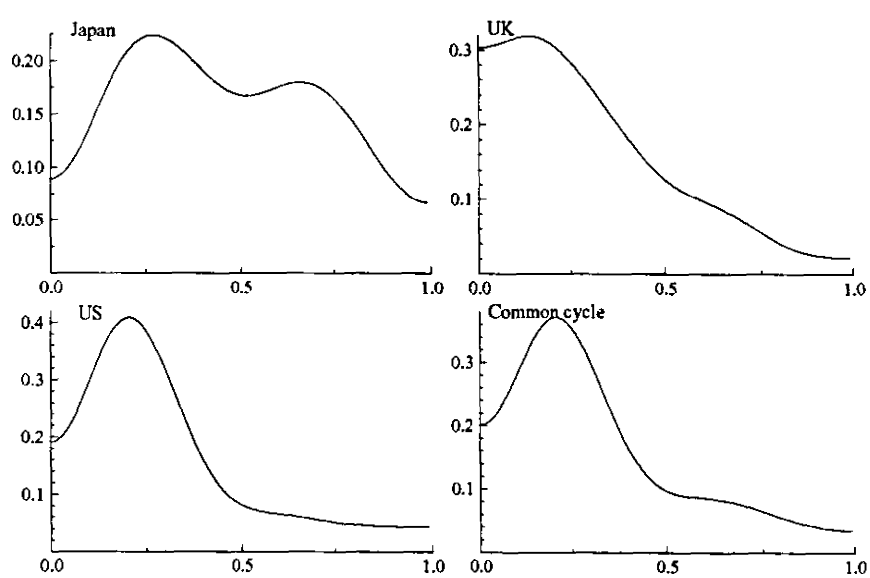

Figure A.26: Spectral densities for the Japan/UK/US group - Output

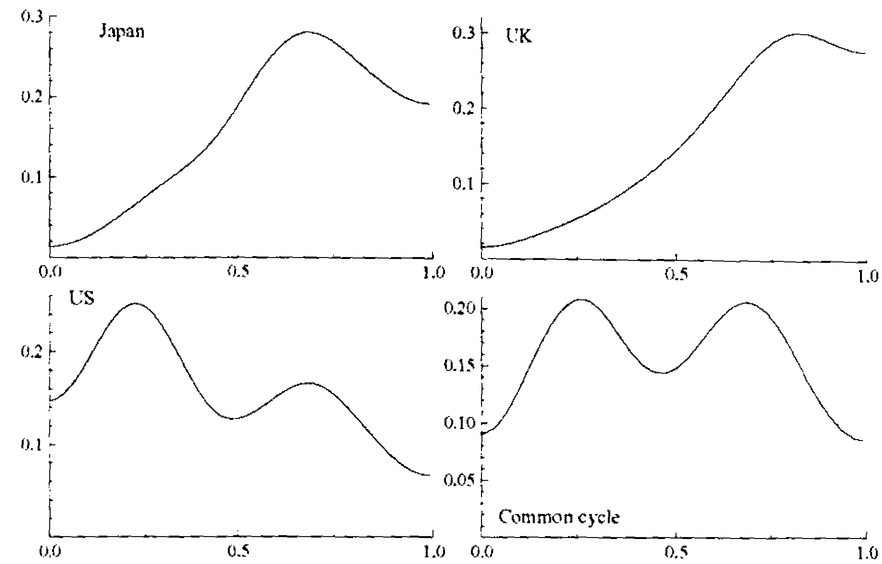

Figure A.27: Spectral densities for Japan/UK/US group- Consumption

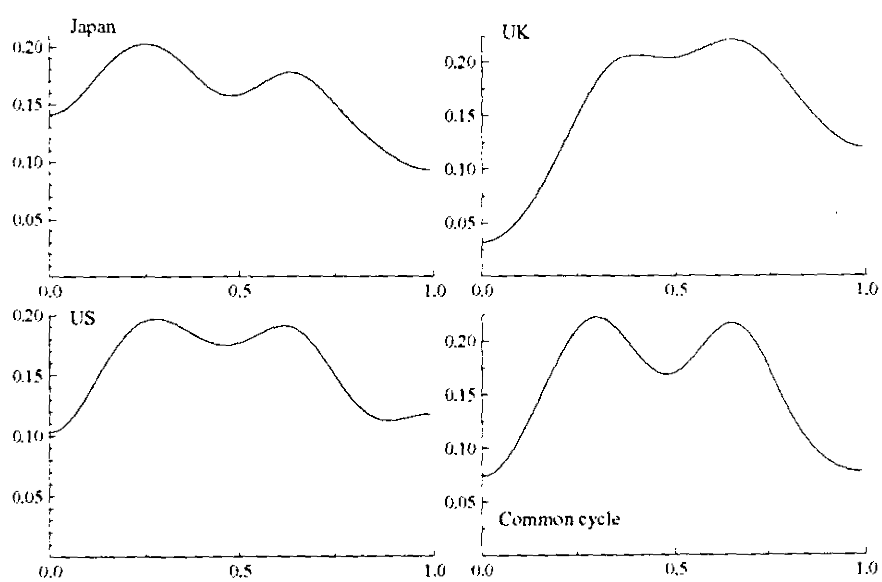

Figure A.28: Spectral densities for Japan/UK/US group- Public expenditures

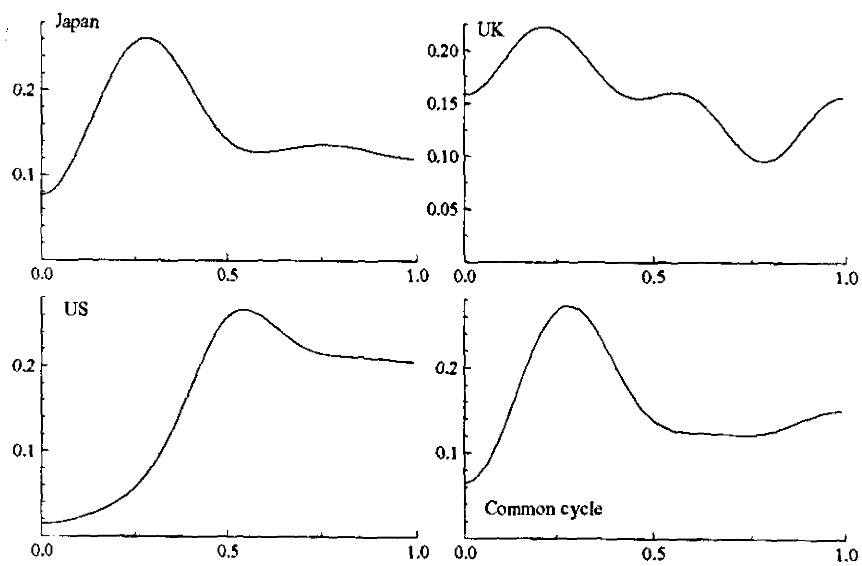

Figure A.29: Spectral densities for Japan/UK/US group- Investment

A.3.7 Simple correlations

Correlations between univariate and common cycles

	Output			Consumption			Public Expenditures			Investment		
	Euro	Euro-UK	UK-Jp-US	Euro	Euro-UK	UK-Jp-US	Euro	Euro-UK	UK-Jp-US	Euro	Euro-UK	UK-Jp-US
Belgium	0.63	0.62		-0.81	0.51		0.08	0.16		0.51	-0.10	
France	0.79	0.79		0.12	-0.22		0.40	-0.68		-0.33	0.50	
Germany	0.34	0.34		0.05	-0.12							
Italy	0.56	0.54		-0.30	0.57		-0.93	0.66		0.67	-0.74	
Netherlands	0.04	0.06		0.43	-0.40		0.05	-0.02		-0.28	0.19	
UK		0.42	0.51		0.60	0.30		0.35	0.51		0.39	0.57
Japan			0.38			0.11			0.72			0.72
US			0.71			0.95			0.65			-0.17
		(0.47			(0.07)			(0.03)			(-0.04)	
Average	0.47	0.46	0.53	-0.10	0.16	0.45	-0.10	0.09	0.63	0.14	0.05	0.37

Table A.7: Simple correlations between individual and common cycles

Correlations between univariate or common cycles

	Output		Consumption		Public expenditures		Investment	
	Euro	Euro+UK	Euro	Euro+UK	Euro	Euro+UK	Euro	Euro+UK
Ja/UK/US	0.17	0.21	-0.02	0.05	0.10	-0.09	0.08	-0.10
UK	0.28	0.42	-0.11	0.60	-0.07	0.35	-0.17	0.39
US	0.15	0.19	-0.06	0.08	0.09	-0.03	0.03	-0.07
Jap.	-0.02	0.00	0.03	-0.16	0.13	-0.05	-0.05	0.06

Table A.8: Simple correlations between individual and common cycles

Correlations - UK/partners

	Output	Cons.	Public exp.	Invest.
Euro	0.28	-0.11	-0.07	-0.17
US	0.20	0.12	0.12	0.05
Jap.	0.03	-0.25	0.13	0.07

Table A.9: Simple correlations between individual and common cycles

A.3.8 Modifications of cycles when the UK is included into the Euro group

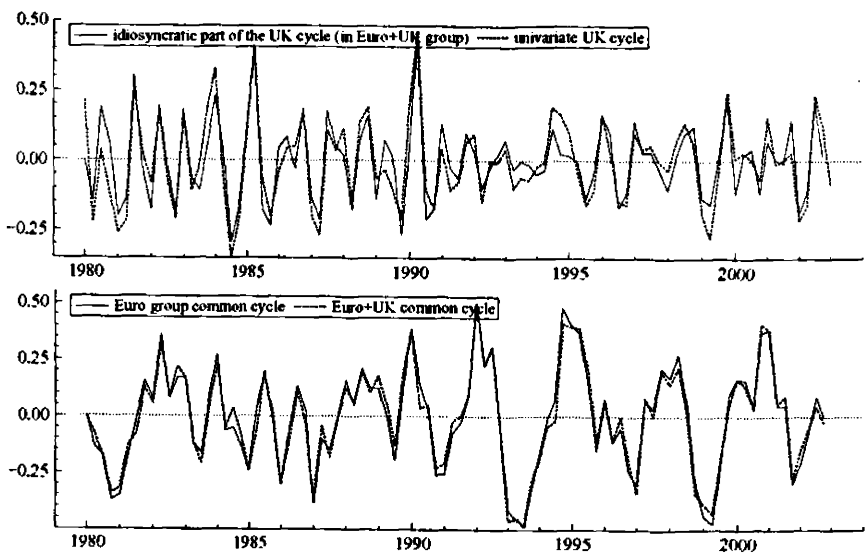

Figure A.30: Modifications of the common cycle and the UK individual cycle induced by the inclusion of the UK into the Euro group – Output

A.4 Appendix for Chapter 4

A.4.1 State-space form

The system to be estimated, eq. (4.5), (4.6), (4.8), (4.2) and (4.3), has to be put under the general state-space form,

$$\alpha_{t+1} = \mathbf{T}\alpha_t + \mathbf{H}\varepsilon_t \tag{A.26}$$

$$\mathbf{y}_t = \mathbf{Z}\alpha_t + \mathbf{x}_t\mathbf{B} + \mathbf{G}\varepsilon_t \tag{A.27}$$

where $\mathbf{Z}, \mathbf{B}, \mathbf{G}, \mathbf{T}$ and $\mathbf{H}$ are matrices of coefficients[21]. eq. (A.26) is the so-called measurement equation and (A.27) the state equation. $\mathbf{y}_t$ is a vector of dependent variables, corresponding to y_t and π_t in our model, α_t is a vector corresponding to the unobserved components, y_t^*, g_t and z_t here. r_t^* is suppressed from the model because it can be fully recovered from g_t and z_t, provided we get c. $\mathbf{x}_t$ is a vector of deterministic variables. $\varepsilon_t \sim NID(0, \mathbf{I})$ and $\alpha_0 \sim N(a, \mathbf{P})$. Note that the representation of (A.28) is particular in that the vector of parameters $\mathbf{B}$ is treated as a vector of unobserved variables. This is a feature of the library *Ssfpack* for *Ox* that we use in this paper.

The system can be rewritten

$$\begin{pmatrix} \alpha_{t+1} \\ \mathbf{B} \\ \mathbf{y}_t \end{pmatrix} = \begin{pmatrix} \mathbf{T} & 0 \\ 0 & \mathbf{I} \\ \mathbf{Z} & \mathbf{x}_t \end{pmatrix} \begin{pmatrix} \alpha_t \\ \mathbf{B} \end{pmatrix} + \begin{pmatrix} u_t \\ 0 \\ v_t \end{pmatrix} \tag{A.28}$$

$$= \Phi \begin{pmatrix} \alpha_t \\ \mathbf{B} \end{pmatrix} + \mathbf{u}_t, \qquad \mathbf{u}_t \sim NID(0, \mathbf{u}_t\mathbf{u}_t')$$

where $\alpha_t' = \begin{pmatrix} y_t^* & y_{t-1}^* & \cdots & y_{t-n}^* & g_t & g_{t-1} & z_t & z_{t-1} \end{pmatrix}$, $\mathbf{B}' = \begin{pmatrix} a_x & \cdots & a_{rx} & b_x & \cdots & b_{sx} \end{pmatrix}$.

[21] We assume here that they are constant since this is the specification used in our model, but they could as well be defined as time varying parameters.

$$
\mathbf{T} = \begin{pmatrix}
1 & 0 & \cdots & 0 & 1 & 0 & 0 & 0 \\
1 & \cdots & 0 & 0 & & 0 & & \\
\vdots & \ddots & \vdots & \vdots & & & \ddots & \vdots \\
0 & \cdots & 1 & 0 & & & & 0 \\
 & & & & 1 & 0 & & \\
\vdots & & & & 1 & 0 & & \\
 & & & & & & 1 & 0 \\
0 & & \cdots & & & 0 & 1 & 0
\end{pmatrix}
$$

$$
\mathbf{Z} = \begin{pmatrix}
1 & -a_{y1} & \cdots & -a_{yn} & 0 & -c.a_r & 0 & -a_r \\
0 & -b_{y1} & \cdots & 0 & 0 & \cdots & \cdots & 0
\end{pmatrix}
$$

$$
\mathbf{x}_t = \begin{pmatrix}
x_{1t} & \cdots & x_{rt} & 0 & \cdots & 0 \\
0 & \cdots & 0 & x'_{1t} & \cdots & x'_{st}
\end{pmatrix}
$$

and $\mathbf{u}'_t = \begin{pmatrix} \varepsilon_{4,t} & 0 & \cdots & 0 & \varepsilon_{5,t} & 0 & \varepsilon_{3,t} & 0 & \cdots & \cdots & 0 & \varepsilon_{1,t} & \varepsilon_{2,t} \end{pmatrix}$.

$$
\mathbf{Y}_t = \begin{pmatrix} Y_1 \\ Y_2 \end{pmatrix} = \begin{pmatrix} y_t - a_{y1}y_{t-1} - \ldots - a_{yn}y_{t-n} - a_r r_{t-1} \\ \pi_t - b_{\pi 1}\pi_{t-1} - \ldots - b_{\pi p}\pi_{t-p} - b_y y_{t-1} \end{pmatrix}
$$

this specification of the vector $\mathbf{Y}_t$ is a simple trick that allows to impose constraints on the coefficients (e.g. y_t and y_t^*). The other exogenous variables, namely the dummies x_t, are estimated as in a classical regression model, i.e. are treated as unobserved variables -upon which no constraint can be imposed.

A.4.2 Sensitivity analysis

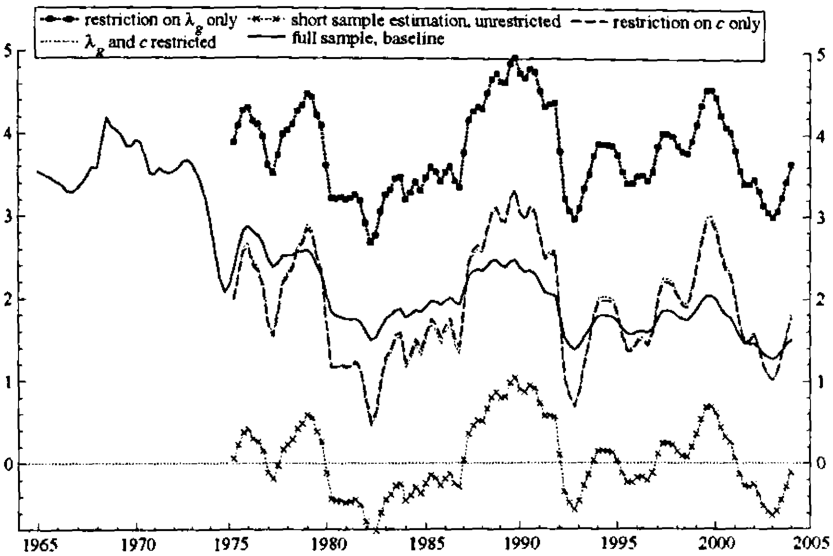

Figure A.31: r_t^* estimates 1973Q2-2004Q2

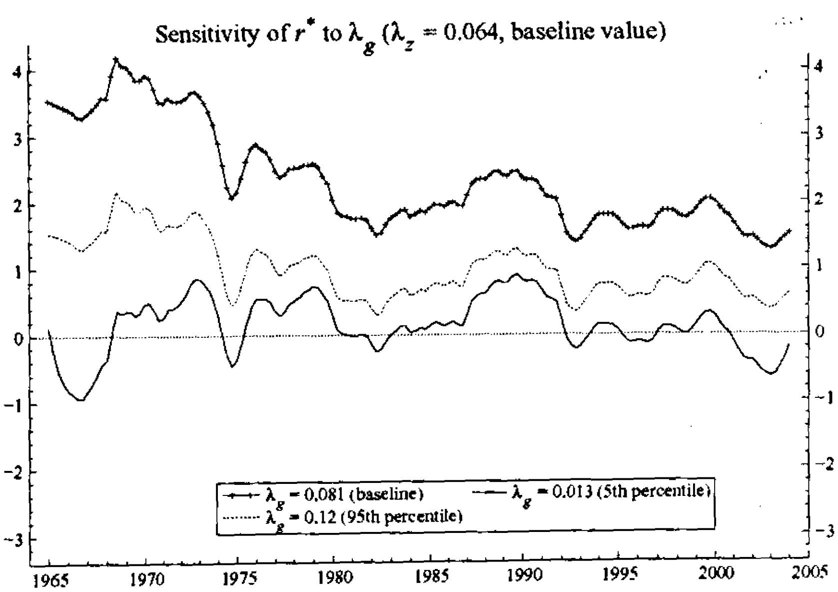

Figure A.32: unrestricted c

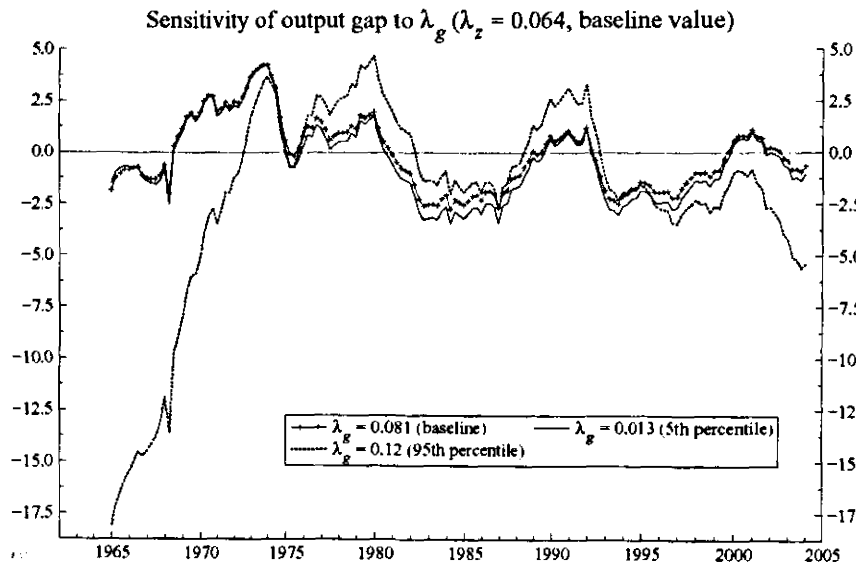

Figure A.33: unrestricted c

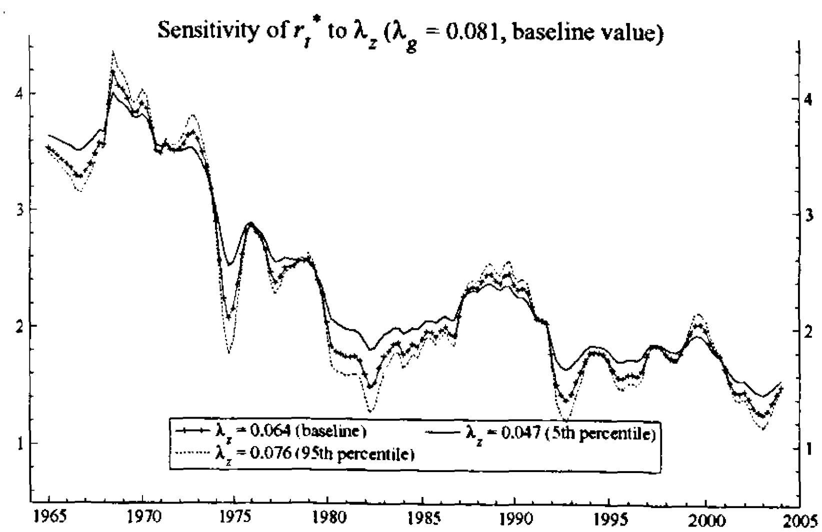

Figure A.34: unrestricted c

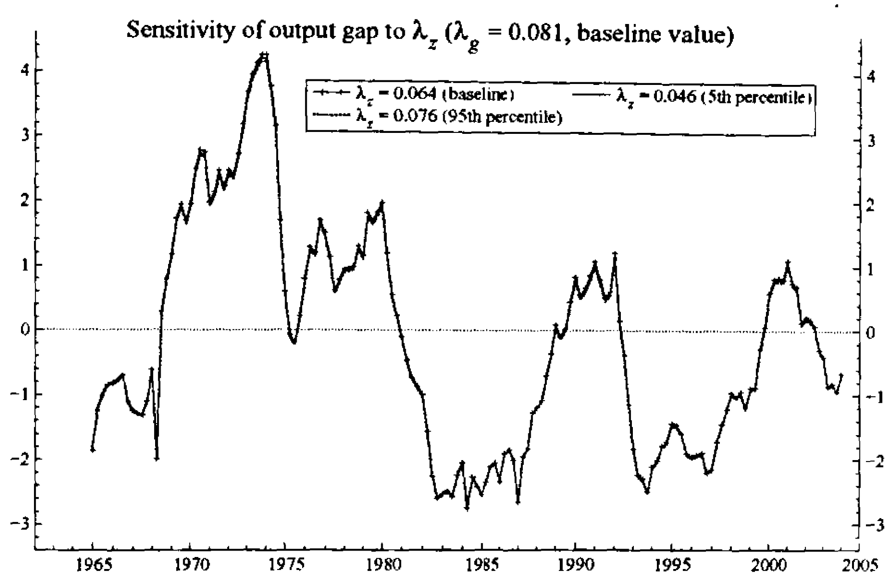

Figure A.35: unrestricted c

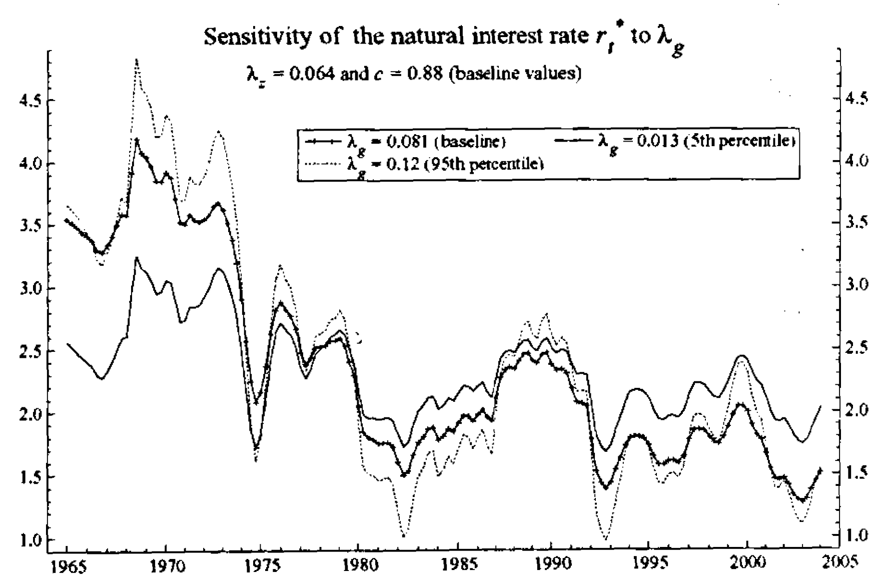

Figure A.36: c restricted to baseline value

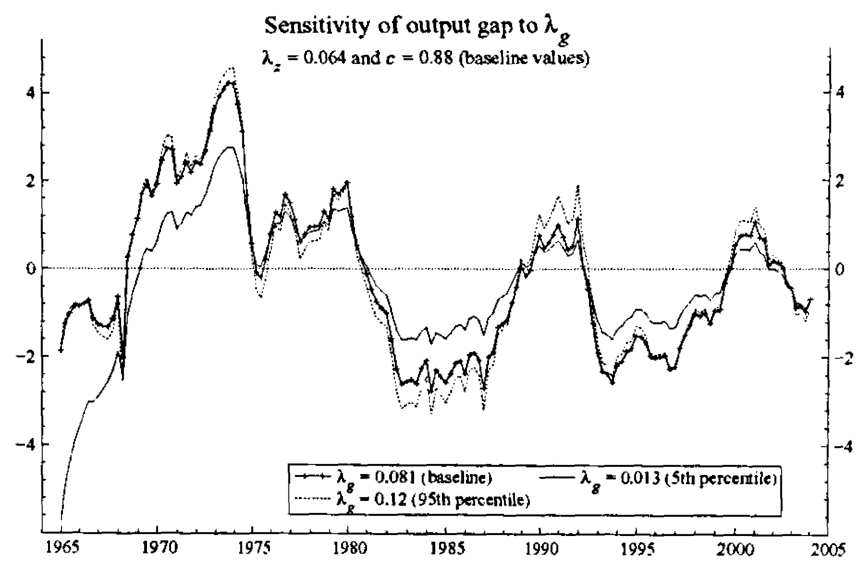

Figure A.37: c restricted to baseline value

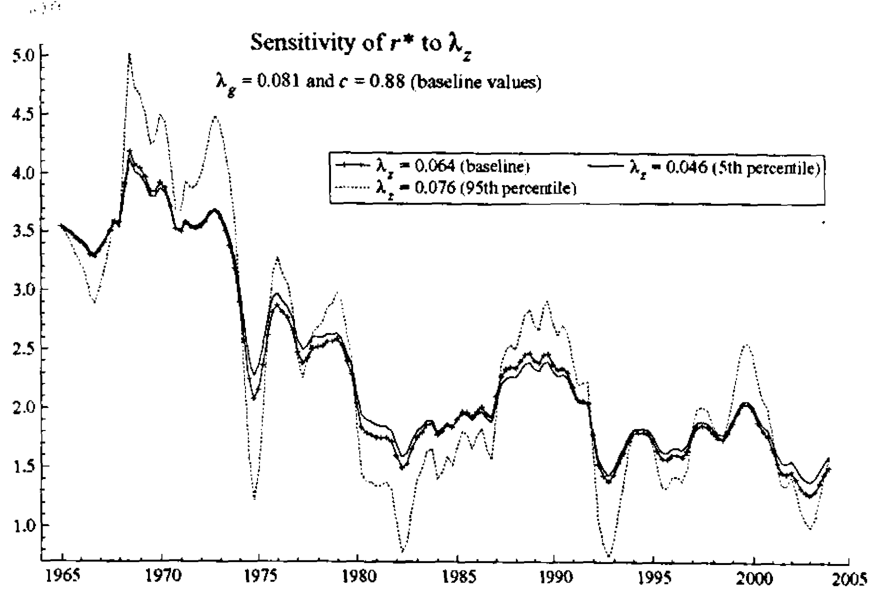

Figure A.38: c restricted to baseline value

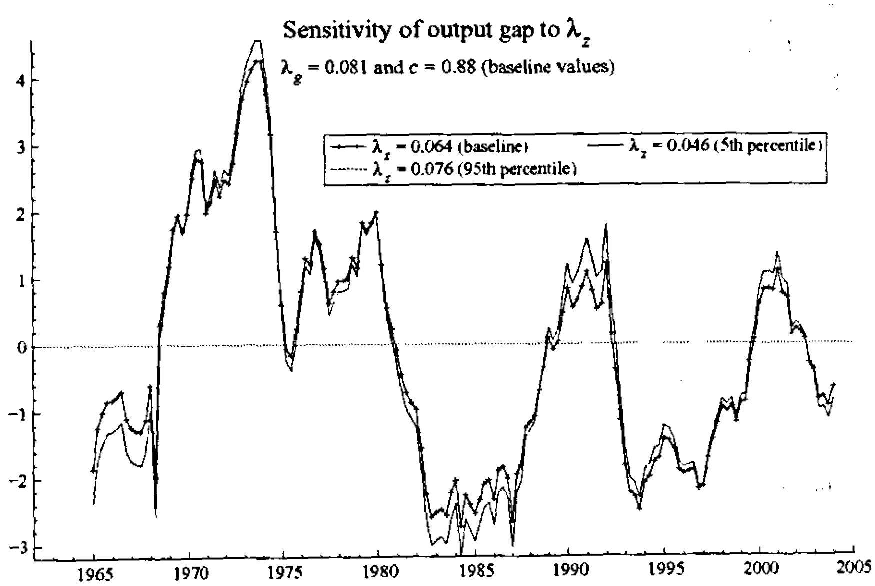

Figure A.39: c restricted to baseline value

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